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Explanatory notes

The following symbols are used in tables in the Review:

… Three dots indicate that data are not available or are not separately reported.

(–) A dash indicates that the amount is nil or negligible.

A blank space in a table means that the item in question is not applicable.

(-) A minus sign indicates a deficit or decrease, unless otherwise specified.

(.) A point is used to indicate decimals.

(/) A slash indicates a crop year or fiscal year; e.g., 2006/2007.

(-) Use of a hyphen between years (e.g., 2006-2007) indicates reference to the complete period considered, including the beginning and end years.

The word "tons" means metric tons and the word "dollars" means United States dollars, unless otherwise stated. References to annual rates of growth or variation signify compound annual rates. Individual figures and percentages in tables do not necessarily add up to the corresponding totals because of rounding.

Creditor protection, information sharing and credit for small and medium-sized enterprises: cross-country evidence

Arturo Galindo and Alejandro Micco

Abstract

Using World Business Environment Survey results for firms in 61 countries, together with country dummies that allow us to deal with observed and unobserved country-specific components, as well as with partial endogeneity, we explore the roles played by creditor protection (e.g. the enforcement of credit contracts) and by the development of credit information mechanisms, such as credit registries, in determining the availability of bank credit for small and medium-sized enterprises (SMEs). We find that better creditor protection and the development of information-sharing mechanisms narrow the financing gap between small and large firms. Countries with poor creditor protection can offset this shortcoming by implementing credit information mechanisms.

I.   Introduction

A large body of literature emphasizes the positive influence that the development of a country's financial sector has on the level and rate of growth of its per capita income. Credit supplied by the banking sector is the most important funding source for firms and households in most countries, in particular developing ones. Unfortunately, bank credit is more expensive and in more limited supply for small and medium-sized enterprises (SMEs) than it is for large firms, which also have a wider range of other financial options. Using comparable cross-section country data on commercial bank credit from 61 countries, while controlling for country fixed effects, we look at the importance of: (i) the role of the legal system in requiring and enforcing loan repayment and the superiority of the common law system in this regard; and (ii) the different types of institutional protection, such as credit bureaux, which can correct the information asymmetries that deter lenders, and their role in shaping access to bank credit. We find that the share of total financing this is obtained from banks is around 30 percentage points greater for SMEs in countries with common law regimes than in those with civil law systems. The presence of a credit registry also increases the level of bank credit for SMEs relative to large firms by around 30 percentage points. Furthermore, our results suggest that creditor rights and the presence of a credit registry are substitutes for one another. Consequently, credit registries have a strong impact on credit in countries with low levels of legal protection (see table 3).

In financial markets, information asymmetries and a lack of contract enforceability constitute a critical barrier to access to finance. However, the financial structure is not independent of firm size. In fact, firm size is a key variable in the analysis of financial restrictions (Beck, Demirgüç-Kunt and Maksimovic, 2005). Large and small firms do not have equal opportunities to access external sources of finance.

Problems of agency costs, information asymmetries and fixed transaction costs result in financial market imperfections. The firms that are typically the most affected by these imperfections are small firms, as their internal information can be rather opaque or, at the least, is not as publicly available as is the case in their larger counterparts. Small firms seeking small loans face higher transaction costs and higher risk premiums because their financial status is less transparent and they have less collateral to offer (Beck and Demirgüç-Kunt, 2006). Similar results have been found by Beck, Demirgüç-Kunt and Maksimovic (2005), Beck and others (2006), and Schiffer and Weder (2001). The latter study also confirms that small firms are confronted with higher barriers to growth.

Recent corporate finance literature has emphasized the role played by financial development and legal protection for outside creditors in terms of corporate performance. One of the crucial implications of these findings is that underdeveloped financial systems may constrain the ability of firms to invest. In these studies, there are two broad views regarding credit to the private sector across economies (Djankov, McLiesh and Shleifer, 2007). The first one stresses the power of creditors. When banks have the ability to force repayment, confiscate collateral or even gain control of the firm, they provide more credit (Townsend, 1979). The second view stresses the importance of information in lending activity. There is a large body of literature that explores the problems posed by moral hazard and adverse selection in financing activity and explains how they reduce access to credit.[1] The more that lenders know about borrowers, their credit history and their total level of debt, the more these problems are mitigated.

The literature on credit markets has identified different ways in which a lender can overcome repayment issues; the most notable of them is the use of collateral. The provision of collateral can serve as a mechanism for mitigating information asymmetries and thus solving credit rationing problems. If the interest rate does not fulfil its role efficiently due to indirect effects on the average quality of credit

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[1]   See Stiglitz and Weiss (1981) and Freixas and Rochet (2008) for a description of the literature.

portfolios, banks can rely on collateral. By taking into account the possibility of pairing collateral with a given interest rate, equilibrium can be achieved without credit rationing.

However, not all loans can readily be backed up with collateral. The collateralization of loans is often problematic for certain types of firms (i.e. new firms, micro-entrepreneurs and SMEs) that often do not have enough fixed assets to offer as collateral. Collateralization is also problematic in countries in which creditor rights are not well protected and where the process of seizing collateral is costly and takes a long time. In this context, the institutional framework for the legal protection of creditors has a particularly strong influence on access to credit, especially for SMEs.

For a review of the importance of collateral in terms of credit contracts and the institutions that support the use of collateral, see La Porta and others (1997 and 1998) and Schiantarelli (1996). Two main findings emerge from this literature: (i) external finance is more costly than internal finance unless loans are fully collateralized; and (ii) the premium on external finance is an inverse function of a borrower's net worth (liquid assets plus the collateral value of illiquid assets). Credit bureaux and/or credit registries can also alleviate the asymmetry of information between lenders and borrowers.[2] The loan and payment history of borrowers reduces information asymmetries and allows lenders to determine their clients' repayment potential. This mechanism creates a different form of collateral —reputation collateral— that banks can use to screen potential borrowers.[3] Credit reporting can therefore be an extremely valuable source of enhanced, fact-based credit risk assessment for creditors and, as such, can facilitate access to financing for SMEs as well as other enterprises. However, the difference lies in the fact that, while credit data and other relevant information on large corporations are generally broadly available, this is usually not the case of information on SMEs.

Facilitating the flow of such credit-related data and other relevant financial information on SMEs helps to alleviate their financing constraints. Ensuring that creditors have ready access to accurate, meaningful and sufficient information on SMEs in a systematic and timely manner enhances their ability to assess SME creditworthiness and hence increases SME access to credit financing. Using a survey on banks, Beck, Demirgüç-Kunt and Martínez Pería (2008) confirm the importance of the availability of sufficient credit history data on SMEs: 70% of developing-country banks and 44% of developed-country banks that responded to the survey stated that the existence of a credit bureau in their country facilitated SME lending.

The purpose of this study is to analyse the importance of creditors' legal rights and borrowers' information-sharing mechanisms in accounting for the variations in access to bank credit of firms of different sizes across countries. In the presence of inefficient legal protection and a lack of borrower information, banks have to monitor their borrowers more closely. Given fixed monitoring costs, lending to SMEs becomes less profitable vis-à-vis lending to large firms and, therefore, the lack of legal protection and information can be expected to disproportionately increase financial restrictions for SMEs.

Most cross-section studies fail to control for the possible endogeneity of creditor protection and information institutions (credit bureaux). It is likely, for example, that, following an increase in bank lending, countries will tend to develop stricter creditor protection laws and credit registries. Consequently, a simple cross-country regression between bank lending and creditor protection/credit bureaux is likely to overestimate the causal effect of these two factors on corporate lending.

To avoid this endogeneity problem, we used dummy countries to run a difference-in-difference Tobit model at the firm level to compare bank credit to SMEs with bank credit to large firms across

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[2] Credit registries may be private or public, and their practices vary in terms of the information that they collect, whether on positive or negative credit behaviour, or both.

[3] See Pagano and Jappelli (1993), Padilla and Pagano (1997) and Bennardo, Pagano and Piccolo (2009) for theoretical papers on the alleviation of information asymmetry.

countries with different levels of legal protection and information institutions. This methodology allows us to control for any country-omitted variable and for the endogeneity problem just described. To assess the importance of creditors' legal rights, we use the "creditor rights" index first proposed by La Porta and others (1997 and 1998). The index measures the legal rights of creditors vis-à-vis defaulting debtors in different jurisdictions. To assess the importance of information in terms of credit, we use the presence and coverage of credit bureaux. The role of credit bureaux is to collect, distribute and, in many cases, analyse information on a borrower's behaviour that has been gleaned from a variety of sources in order to enable creditors to screen potential clients. As noted earlier, the information compiled by private or public credit registries, whether on positive or negative credit behaviour, or both, varies.

We find that the development of credit information mechanisms —as measured, among other metrics, by the population coverage of credit bureaux or public credit registries— is an important factor in narrowing the gap between small and large firms in terms of their access to bank finance. A one-standard-deviation increase in the coverage of credit-reporting institutions reduces the financing gap for small firms by nearly half. In countries where credit registries cover less than 1% of the total population, the share of investment in small firms financed by bank credit corresponds to only 40% of the share of investment in large enterprises that is financed by banks. This figure rises to 63% in countries where credit registries cover more than 1% of the total population (see figure 1A).

We also find that the extent of credit constraints on smaller firms depends on the quality of the regulatory framework, suggesting that, in countries where creditor rights are protected (and enforced), smaller firms have greater access to bank credit for investment purposes. In our sample, this effect is large. In common-law countries (where creditor protection is high), the share of investment financed with bank credit in small firms corresponds to 75% of the share financed in large enterprises. This percentage falls to 50% in non-common-law countries (see figure 1B).

The rest of this study is organized as follows. In section II, we briefly describe the available empirical literature on the impact of credit bureaux and legal protection on credit markets. Section III describes the econometric methodology that has been used. Section IV presents the data and some unconditional results. Sections V and VI present our baselines and robustness results. Section VII concludes our findings.

II.   A review of the empirical evidence

Empirical evidence at the aggregate country level supports the idea that enforcing creditors' legal rights and information-sharing via credit bureaux or public credit registries have a positive impact on credit markets in terms of access to credit, interest rates and default rates.[4]

Using country-level data on 129 different jurisdictions for 1978-2003, Djankov, McLiesh and Shleifer (2007) find that both creditor protection and information sharing have a positive correlation with credit relative to GDP. Although both types of institutions play a complementary role in fostering private credit, these authors find that the effectiveness of each of these institutions varies across countries, depending on the type of legal system that is in place. While legal protection of creditors is associated with common law traditions, credit bureaux and public credit registries are more effective in civil-law countries.

Jappelli and Pagano (2002) provide evidence similar to the evidence presented by Djankov, McLiesh and Shleifer (2007), although for a much smaller sample of countries. Similarly, in a cross-

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[4]   Galindo and Micco (2007) present evidence that financial development also has an impact on credit volatility; Feldmann (2013) provides evidence which indicates that financial development has an impact on firms' labour decisions as well.

country study, Warnock and Warnock (2008) show that the development of mortgage credit markets is positively correlated with the development of credit bureaux.[5]

Empirical evidence at the firm level is scarcer but is necessary in order to assess the impact of information sharing on credit access that is conditional on the characteristics of the firms that are seeking loans. Galindo and Miller (2001) analysed cross-sectional balance sheet data (mostly from large, listed firms) and found that information sharing reduces credit constraints, particularly for small and young firms. They estimated investment equations and found that investment was less sensitive to a firm's cash flow —a metric traditionally used to assess credit constraints— in countries with more developed information-sharing institutions. As mentioned earlier, these cross-country estimates are likely to be upward biased due to endogeneity.

Love and Mylenko (2003) use World Business Environment Survey data to test the impact of the presence of a credit bureau on the perception of firms facing credit constraints and on an increased probability of a firm relying on bank lending. They find that the existence of a private credit bureau is associated with situations in which few firms report that they are financially constrained and more companies report that they rely on bank credit. This last of these results is stronger for small and medium-sized firms.

There is little evidence at the firm level regarding the impact and interaction of the presence of credit bureaux and the existence of legal creditor protection policies. Brown, Jappelli and Pagano (2009) use cross-sectional estimates and a panel of information on transition countries in Eastern Europe to assess the role of information sharing in countries with weak corporate laws and creditor rights. They find that, on aggregate, information sharing is associated with more abundant and cheaper credit. At the firm level, based on cross-sectional data, they find that information sharing and transparency are substitutes for one another in improving credit access. This is the first study to include panel data in an assessment of the impact of information sharing, but the results are inconclusive regarding the relationship between credit bureaux, creditors' legal protection and access to credit for different types of firms (in terms of size and transparency). This is partly because their sample does not provide cross-country variation with regard to creditors' legal protection. In addition, these authors do not control for countries' individual time trends during a period when transition economies were undergoing many different types of changes.

We use a difference-in-difference approach to analyse the impact of creditor protection and information sharing on the share of investment financed with bank credit. Using a similar approach and cross-sectional data from the Business Environment Surveys of the World Bank, Galindo and Micco (2005) show that weak legal protection has a stronger negative impact on SMEs.[6] In line with Djankov, McLiesh and Shleifer (2007), they find that the gap in credit access between SMEs and large firms is bigger in non-common-law countries.

None of these previous studies looked at the impact or interaction of credit bureaux and legal creditor protection policies using a difference-in-difference methodology.

III. Econometric methods

Our hypothesis implies that, compared to large firms, SMEs should have greater relative access to credit in countries where the legal rights of creditors are strong and information sharing is more developed. The dependent variable in our study is the share of financing that comes from banks. Given

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[5] Other authors emphasize the impact of credit bureaux on private credit and find that credit reporting helps lenders to reduce default rates. Examples include IDB (2004), Powell and others (2004), Barron and Staten (2003) and Kallberg and Udell (2003).

[6] This is an earlier working paper which focuses only on creditor protection.

that this variable is restricted to between 0 and 100, we use the Tobit censored regression model and introduce dummy countries to capture country fixed effects.[7] Although Greene (2004) shows that the location coefficients in the Tobit model, unlike those in the probit and logit models, are unaffected by the "incidental parameters problem", we use ordinary least squares (OLS) estimations for robustness. In order to control for relevant firm-level characteristics that may affect access to bank credit, we estimate empirical models at the firm level. For that purpose, we control for variables commonly used in the literature, such as the firm's export orientation or lack thereof, the firm's ownership structure (whether it is government-owned or foreign-owned), and use sector dummies which indicate the area in which the firm operates. To capture the difference in the financing gap associated with different levels of information sharing and creditor protection, we include size dummies (Small and Medium) and interactions between these dummies and our measures of information sharing and creditor rights.

We exploit country/firm variations by estimating the following Tobit model:

$$C_{ijc} = \begin{cases} C^*_{ijc} & if \quad 0 < C^*_{ijc} < 1 \\ 0 & if \quad C^*_{ijc} \leq 0 \\ 100 & if \quad C^*_{ijc} \geq 1 \end{cases}$$

Where C* is the following latent variable:

$$C_{ijc} = \alpha_j + \alpha_c + \delta X_{ijc} + \delta_1 Small_{ijc} + \delta_2 Medium_{ijc} + \delta_3 Small_{ijc} * Inf_c$$
$$+ \delta_4 Medium_{ijc} * Inf_c + \delta_5 Small_{ijc} * CR_c + \delta_6 Medium_{ijc} * CR_c + \varepsilon_{ijc}$$

where C_{jjc} denotes the share of investment financed with bank credit in firm i, sector j and country c; α_j and α_c are sector and country fixed effects; X_{ijc} is a vector of firm-specific variables; $Small_{ijc}$ is a dummy variable equivalent to 1 if the firm is small; Inf_c is a measure of information sharing in country c; CR_c is a measure of creditors' legal protection, and ε_{ijc} is a normal error term. In some specifications, we include interaction terms between Inf_c and CR_c.

We expect a negative coefficient for the Small and Medium dummies, since small and medium-sized firms have less access to bank credit than large firms (our control group). Our main variables of interest are size dummies interacting with CR and Inf. We expect a positive sign for each of these four variables. Improvements in creditor rights or the introduction of credit registries (better information) should increase access to bank credit for all firms (main effect) but in particular for SMEs (additional effect).

We control for country-level fixed effects to capture any institutional or macroeconomic variable that may also affect access to bank credit. In particular, the country dummies account for the main effect of stronger legal rights for creditors and information sharing. Given that the size dummies interact with variables that do not vary at the country level, we use clustered standard errors to adjust them.[8] This is extremely important, since the variables that interact with the size variables do not vary at the firm level but only at the country level. Moulton (1990) demonstrated the serious downward biases in the estimated standard errors that can result from estimating the effects of aggregate explanatory variables on individual-specific (firm-specific, in this case) response variables. Clustered standard errors help to reduce that bias. We weight observations by the inverse of the number of firms in each country-size cell to control for the different number of firms across countries.[9] We also control for sector-specific effects.

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[7] The Hausman-McFadden specification test for IIA rejects the null hypothesis that coefficients for the model with and without country dummies are the same (chi(13)=78.6).

[8] In particular, we cluster at the country-size level. See Moulton (1990) and Judson and Owen (1996).

[9] Without weights, countries with more observations will drive the results, although our variables of interest vary only across countries.

The choice of our empirical methodology is closely related to recent research by Greene on fixed effects in limited dependent variable models. The authors of many firm-level studies have opted to use random-effect Tobit models to estimate the impact of country-wide variables on firm-specific truncated indicators, such as the share of investment financed by credit, in which accounting for individual effects appears relevant.[10] However, Greene (2002 and 2004) shows that, if the explanatory variables are not uncorrelated with the individual effects (a usually unpalatable assumption), the random-effect model can lead to biased estimates of the model's slope parameters. In such a case, the fixed-effect Tobit is a preferable methodology, given that the bias in the slope parameters attributed to the incidental parameter problem tends to be negligible.

Country-level fixed effects allow us to deal with observed and unobserved country-specific components, as well as, in some cases, partial endogeneity and inverse causality. Any increase in total credit that induces the development of credit bureaux and creditor protection mechanisms is controlled for by the fixed effect. Only pathological changes in the relative amount of credit to SMEs that imply changes in credit registries and creditor rights at the country level will not be taken into account by the country fixed effect; thus, some scope for reverse causality will remain.

IV. Data

This section describes the data sources and variables used in the empirical analysis. Our main source of data is the World Business Environment Survey (WBES);[11] other sources include several research papers that provide valuable information on the development of credit bureaux and creditor protection regulations around the world.[12] For the purposes of this study, the dependent variable is the leverage of firms of different sizes. Our purpose is to test if access to credit, defined as the share of investment financed with bank credit, depends on the extent of creditors' legal protection, the development of information-sharing mechanisms, the size of firms and/or the interaction between size, on the one hand, and creditor protection and information sharing, on the other.

WBES results provide a firm-level dataset that consists of responses from more than 10,000 firms across the world to questions related to a country's business environment. The survey includes questions regarding firms' financing structure. Business managers were asked to report how much of their investment was financed over the last year from the following sources: (i) retained earnings, (ii) funds from family and friends, (iii) equity, (iv) supplier credit, (v) leasing arrangements, (vi) money lenders, (vii) other public-sector support, (viii) local commercial banks, (ix) foreign banks, (x) development banks, and (xi) other. For our purposes, we define the dependent variable as the sum of the portion of investment financed with credit provided by local commercial banks and foreign banks and label this as "access to bank credit."

When constructing the "access to bank credit" variable, we took great care to exclude erroneous data. We dropped all firms that reported percentages for funding sources that totalled less than 90% or more than 110%. Thus, we allowed for the possibility of small mistakes in addition, but eliminated excessively erroneous data.

Another crucial firm-level variable in our analysis is the size of firms. These data were also obtained from WBES, which classifies firms into three different groups by size: small firms are defined as those with more than 5 but fewer than 50 workers; medium-sized firms are those with more than 50

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[10] See, for example, Beck, Demirgüç-Kunt and Maksimovic (2001).

[11] This dataset has been used in various cross-country studies. See, for example, Beck, Demirgüç-Kunt and Maksimovic (2005) and Beck and others (2006).

[12]   See, for example, La Porta and others (1997 and 1998) and Djankov, McLiesh and Shleifer (2007).

but fewer than 500 employees; and large firms are those with more than 500 workers. Other firm-level variables included in our empirical analysis that can affect access to finance are the ownership structure of the firm (foreign- or State-owned), export orientation and the economic sector in which the firm operates.[13]

Table 1 reports some basic descriptive statistics for the dependent variable in our study for the 61 countries in which creditor protection laws and credit bureau data can be matched up with WBES data. The average firm in our sample finances 16.2% of its investment with bank credit. As expected, large firms finance a bigger share of investment with credit than medium-sized and small ones do, since large firms face lower information asymmetries and have more collateral, and they therefore find it easier to gain access to credit markets. Export firms, defined as firms for which exports represent at least 1% of their total sales, have greater access to credit than other firms. This may be due to the fact that export activity signals high productivity and thus high repayment probabilities to creditors and hence eases financial constraints. This would hold true for firms of all sizes. In the sample there are no relevant unconditional differences between the shares of investment financed by small and medium-sized foreign-owned firms.

Table 1
Share of investment financed with bank credit: descriptive statistics
(In levels and percentages)

		All	Export binary	Foreign binary	State-owned firms
All[a] 6 616	Mean	16.2	23.2	11.6	23.3
	Standard deviation	27.0	30.4	24.1	31.7
Small 2 803	Mean	11.3	17.5	12.1	14.2
	Standard deviation	22.8	26.8	27.4	26.1
Medium 2 797	Mean	17.5	22.4	9.3	20.8
	Standard deviation	27.7	29.9	21.8	30.5
Large 1 014	Mean	26.1	29.1	17.2	32.0
	Standard deviation	32.0	32.9	27.5	34.2

Source: Prepared by the authors, on the basis of World Bank, World Business Environment Survey (WBES).

ª   Not all firms record this variable, so the total does not correspond to the sum of firms by size.

To measure the development of credit registries, we use data from the Doing Business project of the World Bank. Our main variables are the presence of any credit registry, a dummy variable equivalent to 1 if there is a public credit registry or a private credit bureau, and the coverage of private and public registries. Private credit bureau coverage reports the number of individuals or firms listed by a private or public credit bureau with information on repayment history, unpaid debts or credit outstanding. The number is expressed as a percentage of the adult population. We construct an index taking the larger of the two as the relevant variable for the economy, and we also explore the role of each type of coverage individually. For example, if a country has a private credit bureau that covers 550 per 1,000 of the population and a public credit registry that covers 200 per 1,000 of the population only, we use the larger of the two (550). The reason for doing this is to avoid double counting borrowers (many of whom may appear in both registries). In a non-reported exercise, we use each of the coverage measures separately. This measure gives us some insights into how many SMEs may be covered by credit registries. For example, for the case of Chile, a mid-developed country in our sample, the private credit registry has 227 entries per 1,000 adults. For that same year, Chile had

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[13] The dataset does not include the number of employees, sales figures or the level of assets. This study therefore uses the only size variable available.

1.9 large firms per 1,000 adults, 1.5 medium-sized firms per 1,000 adults, and 25 small firms and 123 micro-enterprises per 1,000 adults.[14] The annex reports the countrywide data used in the study.

To proxy the status of creditors' legal protection, we use a set of variables that has frequently been cited in related literature.[15] These variables are measures of certain types of institutions and of rules and regulations that directly affect creditors' ability to seize collateral effectively and efficiently.[16] Following Galindo and Micco (2007), we use a measurement of effective creditor rights that combines a legal variable that reflects creditor protection based on La Porta and others (1997) and an indicator of the rule of law (derived from the World Bank). This variable captures not only the regulatory framework surrounding the rights of creditors in bankruptcy proceedings but also the extent of enforcement of bankruptcy laws. We also proxy the protection of creditor rights using the type of legal system that is in place.[17] We draw here on the proposition espoused by La Porta and others (1997) that countries with common law systems are characterized by better creditor protection.

Figure 1 and table 2 report some basic statistics that underlie the econometric study. Figure 1 shows the average share of investment that has been financed with bank credit for firms of different sizes in countries in which the development of credit bureaux has reached differing levels (see figure 1A)[18] and which have differing legal systems (see figure 1B).

The development of credit bureaux is measured as the portion of the population that is covered either by a private credit bureau or a public credit registry. We split the sample between countries with coverage higher or lower than 1%, with the aim of having a similar number of countries in each group. As expected, larger firms are able to access more credit to finance their investment. However, in countries with stronger credit registries, firms of all sizes have more available credit than they do in countries in which the presence of credit bureaux is weaker. The difference in access to credit is proportionally greater for small firms. We find similar results when we split the sample between common-law and non-common-law countries. In both groups, large firms have more access to credit than small firms do, but this difference is fairly small in common-law countries.

In addition, in table 2, the sample is divided between countries with high and low coverage indices for credit registries or credit bureaux, using as a threshold a coverage level of 1% of the total population (10%); the sample is also divided according to the level of protection of creditors' rights. As expected, in each quadrant, the share of investment financed with bank credit increases with the size of the firms involved. Also as expected, the share of finance is larger in countries with stronger creditor protection for any given segment of firm size. What is particularly interesting to note is how the share of credit in small firms reaches a similar level in countries with strong creditor protection and no credit registry and in countries with weak creditor protection and a credit registry. The presence of a credit bureau or a credit registry therefore seems to compensate for deficiencies in creditor rights regulations. This finding is important for developing countries, since it appears that a weak institutional set-up in the financial sector, which could be hard to rectify in the short term, may be compensated for by a greater supply of information.

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[14] In Chile, large, medium-sized and small firms are defined as those having annual sales above US$ 3.8 million, US$ 961,000, and US$ 92,000, respectively.

[15] See La Porta and others (1997 and 1998) and Galindo and Micco (2007).

[16] The creditor rights index measures: (i) whether there are restrictions, such as creditor consent, when a debtor files for reorganization; (ii) whether secured creditors are able to seize their collateral after the petition for reorganization is approved (in other words, whether there is no "automatic stay" or court-ordered "asset freeze"); (iii) whether secured creditors are paid first out of the proceeds of the liquidation of a bankrupt firm, and (iv) whether an administrator, as opposed to management, is responsible for running the business during the reorganization.

[17] Several authors have linked a common-law legal tradition with better protection for creditors. See, for example, La Porta and others (1997 and 1998).

[18] We use a maximum coverage of 1% (10/100) as a cutoff.

Figure 1
Share of investment financed with bank credit, by firm size
(In percentages)
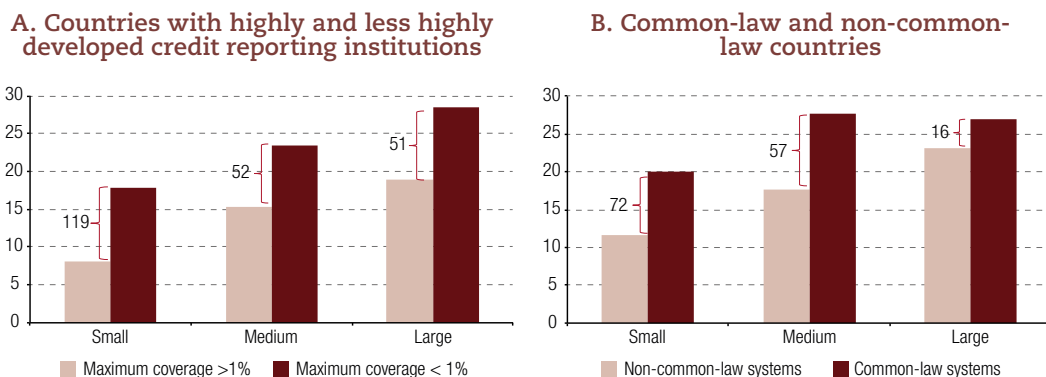
**A. Countries with highly and less highly
developed credit reporting institutions**

**B. Common-law and non-common-
law countries**

Source: Prepared by the authors, on the basis of World Bank, World Business Environment Survey (WBES).
Note: These figures do not include Portugal because it has the same number of observations in each group.

Table 2
Average share of investment financed with bank credit, by firm size, presence or absence
of information-sharing mechanisms and type of legal system
(In percentages)

	Non-common-law countries			Common-law countries		
	Small	Medium	Large	Small	Medium	Large
No private or public registry	5.4	10.3	13.6	25.9	34.4	35.0
Number of countries	18	18	18	4	4	4
Private or public registry	15.2	21.7	28.7	16.8	24.0	22.3
Number of countries	32	32	31	7	7	7

Source: Prepared by the authors, on the basis of World Bank, World Business Environment Survey (WBES).

V.   Baseline results

Our baseline results are reported in table 3. With respect to firm-level controls, we find that exporters finance around 10% more of their investment with bank loans than firms that cater to the domestic market.[19] We do not find significant differences between the shares of financing for State-owned and foreign-owned firms. Finally, although not reported in the tables, firms in the manufacturing sector, perhaps due to the tangibility of their assets, have greater access to bank loans. In a country with an average coverage of credit information (113), small firms finance about 30 percentage points less of their investment with credit than large firms do, while the corresponding differential for medium-sized companies is about 11.5 percentage points.[20]

Focusing on the variables of interest for this study, we find that the development of credit-sharing mechanisms alters the financing gap between large and small firms. Column 1 indicates that a one-standard-deviation increase in the coverage of credit information-sharing institutions (119 per 1,000 inhabitants) above the average (113) reduces the financing gap between large and small

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[19] As a measure of robustness, we use a different definition of export firms (exports > 10% of sales) but the results do not change.

[20] These results come from column 1. To compute the value for small and medium-sized firms, we take CL=0.5 (CL is either 0 or 1) and max coverage = 113 (the average maximum coverage in the sample): -30 = -48 + 0.07 x 113 + 19.9 x 0.5.

firms to nearly 22 percentage points (from 30 points) and the gap between large and medium-sized firms to 8.5 percentage points (from 11.5).[21] The effect of information-sharing mechanisms is not only statistically significant, but also is actually quite large in relation to the prevailing financing gaps between large and small firms.

Column 2 also shows that the presence of credit-sharing institutions is influential in explaining the financing gap between large and small firms. Our dummy variable for the presence of a public registry or a private credit bureau indicates that the presence of a credit registry reduces the gap between large and small firms by 39 percentage points. For medium-sized firms, the gap narrows by 17 percentage points. Column 3 shows qualitatively similar results when dummy variables are used for the presence of information-sharing institutions and their coverage. The presence of credit information institutions significantly increases firms' access to credit, with a differential impact being observed for smaller ones.

Table 3
Baseline econometric results

Dependent variable: share of investment financed by bank credit (firm level)	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Export firms	10.546	10.195	10.191	10.174	9.690	10.248	9.812	3.781
	[2.369]***	[2.366]***	[2.379]***	[2.359]***	[2.320]***	[2.335]***	[2.339]***	[1.058]***
Public ownership (firm)	-6.535	-7.740	-8.106	-7.613	-6.550	-5.634	-6.903	-4.265
	[4.201]	[4.216]*	[4.224]*	[4.224]*	[4.177]	[4.182]	[4.212]	[1.741]**
Foreign ownership (firm)	-0.266	-0.250	-0.039	-0.336	-0.266	-0.462	-0.232	0.116
	[3.080]	[3.075]	[3.071]	[3.076]	[3.067]	[3.093]	[3.064]	[1.456]
Small firms	-48.184	-66.341	-64.858	-68.018	-57.882	-43.239	-58.035	-10.690
	[4.980]***	[9.290]***	[9.295]***	[9.323]***	[8.264]***	[4.794]***	[8.339]***	[1.702]***
Medium-sized firms	-23.681	-32.804	-32.696	-34.610	-27.403	-19.078	-27.365	-3.627
	[4.007]***	[6.859]***	[6.879]***	[7.199]***	[6.080]***	[3.735]***	[6.121]***	[1.346]***
Credit registry coverage	0.070		0.042			0.058	0.028	0.013
Small firms	[0.013]***		[0.017]**			[0.013]***	[0.016]*	[0.005]**
Credit registry coverage	0.025		0.010			0.015	-0.001	0.004
Medium-sized firms	[0.010]**		[0.011]			[0.011]	[0.014]	[0.004]
Private or public registry		39.089	28.189	35.888	33.617		27.299	
Small firms		[8.962]***	[10.500]***	[11.144]***	[8.315]***		[9.401]***	
Private or public registry		17.511	15.440	13.970	14.934		15.243	
Medium-sized firms		[6.817]**	[7.682]**	[7.934]*	[6.464]**		[7.423]**	
Registry with positive information				5.787				
Small firms				[9.737]				
Registry with positive information				6.296				
Medium-sized firms				6.296				
Common-law country	19.878	32.019	26.009	32.643				6.510
Small firms	[7.833]**	[9.042]***	[9.966]***	[8.783]***				[2.623]**
Common-law country	18.764	22.334	21.293	23.023				7.884
Medium-sized firms	[5.407]***	[5.588]***	[5.838]***	[5.675]***				[2.369]***
Effect of creditor rights (country)					26.992	17.790	21.639	
Small firms					[4.522]***	[5.052]***	[5.551]***	

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[21] For small firms, the estimated reduction in the credit gap is 119*0.07, while for medium-sized firms it is 119*0.025.

Table 3 (concluded)

Dependent variable: share of investment financed by bank credit (firm level)	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Effect of creditor rights (country)					17.416	15.430	17.627	
Medium-sized firms					[4.425]***	[4.729]***	[5.099]***	
Observations	6 470	6 604	6 604	6 604	6 470	6 604	6 604	6 604
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Sector fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Number of countries	61	61	61	61	61	61	61	61

Source: World Bank, World Business Environment Survey (WBES) and S. Djankov, C. McLiesh and A. Shleifer, "Private credit in 129 countries", *Journal of Financial Economics*, vol. 84, No. 2, Amsterdam, Elsevier, 2007.

Note: Robust standard errors in brackets. Clusters by country size. * significant at 10%; ** significant at 5%; *** significant at 1%.

Column 4 shows if there is a difference between credit registries which include positive information (a good record of repayment) and credit registries that include only negative information in their credit reports. We do not find any statistically significant difference between the inclusion of positive and negative versus solely negative information in credit reports.

A non-reported regression includes additional interactions that separate the coverage of public credit registries and private credit bureaux. The results suggest that the impact on small and medium-sized firms is not significantly different when discriminating between the two types of institutions, although the impact on medium-sized firms appears to be driven mostly by information coming from private credit bureaux.

To study the role of legal protection for creditors, we include the interaction of the firm-size dummies with measures of credit rights protection. In columns 1 to 4, we interact firm size with a dummy that indicates if a country has a common-law regime, which has been proven to be a good proxy for effective creditor rights (CL). The credit gap between small and large firms is between 20 and 30 percentage points smaller in countries with common-law regimes than in countries with other legal systems. For medium-sized enterprises, the gap with large firms narrows by between 18 and 32 percentage points.

Following Galindo and Micco (2007), in columns 5 to 7, we interact size with an indicator of effective creditors' rights protection based on the interaction outlined by La Porta and others (1997 and 1998) between creditor rights and the rule of law indicator from the Worldwide Governance Indicators dataset.[22] We find very strong results for the impact of creditor rights regulations in reducing the financing gap between small and large firms and between medium-sized and large firms. A one-standard-deviation increase in our measure of effective creditors' rights (0.516) reduces the gap between small and large firms by from 9 to 14 percentage points and the gap between medium-sized and large firms by between 8 and 9 percentage points. In a non-reported regression in which we use the logarithm of the number of days that the justice system takes to enforce a contract as a measure of creditor protection, we find similar results.

Finally, in column 8 we re-do the specification in column 1 using ordinary least squares (OLS). Estimated coefficients have the same expected sign and are significant at conventional levels. Not surprisingly, due to the censoring nature of the data, estimated slopes are smaller in absolute values.

The main policy driver in the reduction of the financing gap is an improvement in the protection of creditors' rights. Nonetheless, and especially for small firms, the effort to strengthen effective creditor rights —a titanic task— can be bolstered by efforts to develop information-sharing mechanisms. These

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[22] See Kaufmann, Kraay and Mastruzzi (2009).

efforts will provide a bigger pay-off if the benefits of information mechanisms are larger in countries with poor creditor protection. To test this hypothesis, columns 1-2 and 4-5 in table 3 are computed in table 4 while including an interaction term between information-sharing and creditor protection. For each of the four specifications, the interaction term is negative for small and medium-sized firms and is statistically significant in all cases for small firms. This implies that the beneficial effect of information-sharing institutions is larger in countries with poor creditor protection. Non-common-law countries can thus offset the absence of effective creditor rights in their credit markets by establishing information-sharing institutions.

Table 4
Econometric results when controlling for the interaction effect between information sharing and creditor protection

Dependent variable: share of investment financed by bank credit (firm level)	(1)	(2)	(3)	(4)
Export firms	10.300	10.000	10.084	9.499
	[2.359]***	[2.339]***	[2.330]***	[2.307]***
Public ownership (firm)	-7.158	-8.763	-5.810	-6.920
	[4.203]*	[4.212]**	[4.177]	[4.165]*
Foreign ownership (firm)	0.037	-0.278	-0.522	-0.263
	[3.076]	[3.057]	[3.083]	[3.059]
Small firms	-51.739	-72.842	-43.978	-57.367
	[5.693]***	[11.168]***	[4.888]***	[8.188]***
Medium-sized firms	-24.723	-34.245	-19.475	-26.579
	[4.551]***	[7.805]***	[3.789]***	[5.927]***
Private or public registry		47.353		33.772
Small firms		[11.224]***		[8.357]***
Private or public registry		19.290		14.535
Medium-sized firms		[8.177]**		[6.368]**
Credit registry coverage	0.092		0.073	
Small firms	[0.018]***		[0.018]***	
Credit registry coverage	0.032		0.024	
Medium-sized firms	[0.015]**		[0.014]*	
Common law (country)	31.556	55.616		
Small firms	[11.397]***	[13.229]***		
Common law (country)	22.476	30.022		
Medium-sized firms	[7.154]***	[10.190]***		
Effective creditor rights (country)			23.826	44.053
Small firms			[6.171]***	[9.157]***
Effective creditor rights (country)			17.629	30.435
Medium-sized firms			[5.607]***	[7.786]***
Private or public registry x common law		-35.180	-0.037	0.000
Small firms		[17.419]**		
Private or public registry x common law		-10.944		
Medium-sized firms		[12.285]		
Credit registry coverage x common law	-0.051			
Small firms	[0.025]**			
Credit registry coverage x common law	-0.017			
Medium-sized firms	[0.019]			
Private or public registry x effective creditor rights				-26.367
Small firms				[10.542]**

Table 4 (concluded)

Dependent variable: share of investment financed by bank credit (firm level)	(1)	(2)	(3)	(4)
Private or public registry x effective creditor rights				-20.811
Medium-sized firms				[9.425]**
Credit registry coverage x effective creditor rights			-0.037	
Small firms			[0.018]**	
Credit registry coverage x effective creditor rights			-0.020	
Medium-sized firms			[0.015]	
Observations	6 604	6 604	6 604	6 604
Country fixed effects	Yes	Yes	Yes	Yes
Sector fixed effects	Yes	Yes	Yes	Yes
Number of countries	61	61	61	61

Source: World Bank, World Business Environment Survey (WBES) and S. Djankov, C. McLiesh and A. Shleifer, "Private credit in 129 countries", *Journal of Financial Economics*, vol. 84, No. 2, Amsterdam, Elsevier, 2007.
Note: Robust standard errors are shown in brackets. Clusters are by country size. * significant at 10%; ** significant at 5%; *** significant at 1%.

In summary, our baseline results suggest that creditors' legal protection and the development of either private or public credit reporting institutions are strongly correlated with access to credit markets, particularly for small firms. In the next section, we control for the level of development and different subsamples in order to gain a greater understanding of the robustness of these findings.

VI. Robustness

A possible driver of our previous results is that the measurements used to capture the development of credit information institutions or legal protection for creditors are proxies for economic development. Most probably, more developed countries have more robust credit-reporting institutions and stronger effective creditor rights (by virtue of a steadfast respect for the rule of law). In order to deal with this possible bias, we control for economic development in two possible ways: first, by using dummy variables that indicate if a country is a low- or middle-income country according to the World Bank classification (columns 1 to 4 in table 5) and, second, by using purchasing power parity (PPP) GDP per capita (columns 5 and 6). As above, we include both measures of credit information development —the maximum coverage variable and the dummy variable— as well as our measure of creditors' legal protection. The results remain qualitatively the same. While there is a reduction in the point estimates of the maximum coverage variable and the credit registry dummy, the results still point in the same direction: credit registries play a significant role in reducing the financing gap between small and large firms. Not surprisingly, effective creditor rights —our proxies for creditor legal protection, which include the rule of law— are weaker when we include controls for levels of income but, in most specifications, they remain significant. In the case of common-law countries, the results are the same as before.

Table 5
Econometric results when controlling for income level

Dependent variable: share of investment financed by bank credit (firm level)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Export firms	9.627	9.959	9.536	9.956	9.707	9.635	10.139
	[2.344]***	[2.349]***	[2.340]***	[2.355]***	[2.337]***	[2.324]***	[2.344]***
Public ownership (firm)	-7.036	-5.403	-6.210	-5.067	-6.517	-5.964	-5.150
	[4.238]*	[4.209]	[4.234]	[4.208]	[4.239]	[4.230]	[4.220]

Table 5 (concluded)

Dependent variable: share of investment financed by bank credit (firm level)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Foreign ownership (firm)	0.256	0.004	-0.026	-0.280	-0.223	-0.283	-0.434
	[3.108]	[3.108]	[3.097]	[3.117]	[3.086]	[3.088]	[3.092]
Small firms	-35.353	-22.886	-34.358	-23.535	-62.474	-56.641	-45.276
	[11.982]***	[10.704]**	[13.548]**	[13.637]*	[9.231]***	[8.552]***	[4.992]***
Medium-sized firms	-24.777	-17.529	-31.952	-24.138	-31.139	-27.550	-21.946
	[9.415]***	[10.071]*	[11.696]***	[12.675]*	[6.692]***	[6.195]***	[4.010]***
Private or public registry	29.737		28.967		31.678	31.142	
Small firms	[8.904]***		[8.706]***		[9.075]***	[8.830]***	
Private or public registry	14.864		15.564		14.716	15.452	
Medium-sized firms	[6.755]**		[6.791]**		[6.768]**	[6.750]**	
Credit registry coverage		0.036		0.043			0.044
Small firms		[0.014]**		[0.015]***			[0.015]***
Credit registry coverage		0.015		0.019			0.010
Medium-sized firms		[0.013]		[0.013]			[0.013]
Common law (country)	29.137	22.942			28.716		20.642
Small firms	[8.062]***	[7.375]***			[8.582]***		[7.238]***
Common law (country)	23.359	21.660			21.671		19.827
Medium-sized firms	[5.958]***	[6.252]***			[5.616]***		[5.557]***
Effective creditor rights (country)			13.902	8.297		19.634	
Small firms			[7.172]*	[7.116]		[8.328]**	
Effective creditor rights (country)			19.508	16.569		19.088	
Medium-sized firms			[6.943]***	[6.734]**		[7.222]***	
Low-income (country)	-47.813	-44.401	-34.878	-34.416			
Small firms	[11.840]***	[13.013]***	[16.606]**	[17.690]*			
Low-income (country)	-14.649	-14.271	6.737	5.266			
Medium-sized firms	[8.295]*	[10.927]	[13.954]	[15.965]			
Middle-income (country)	-23.197	-17.893	-20.339	-15.815			
Small firms	[7.859]***	[9.106]**	[10.496]*	[12.288]			
Middle-income (country)	-4.600	-2.620	4.544	5.679			
Medium-sized firms	[7.110]	[8.569]	[9.690]	[11.266]			
Ln per capita GDP (country)					12.670	6.009	10.659
Small firms					[3.841]***	[5.984]	[4.113]***
Ln per capita GDP (country)					5.381	-1.660	5.826
Medium-sized firms					[2.872]*	[4.909]	[3.700]
Observations	6 604	6 604	6 604	6 604	6 604	6 604	6 604
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Sector fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Number of countries	61	61	61	61	61	61	61

Source: World Bank, World Business Environment Survey (WBES) and S. Djankov, C. McLiesh and A. Shleifer, "Private credit in 129 countries", *Journal of Financial Economics*, vol. 84, No. 2, Amsterdam, Elsevier, 2007.

Note: Robust standard errors are shown in brackets. Clusters are by country size. * significant at 10%; ** significant at 5%; *** significant at 1%. Ln = Logarithmus naturalis.

Table 6 reports estimates for information sharing and creditors' legal protection using the specification from column 2 in table 3 but while dropping one country from our sample at a time. In all cases, our main estimated coefficients are positive and significant at conventional confidence intervals.

Table 6
Econometric results while excluding one country at a time

| Country | Small firms | | Medium-sized firms | | Private or public registry | | | | Common law (country) | | | | Countries |
| | | | | | Small firms | | Medium-sized firms | | Small firms | | Medium-sized firms | | |
	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	
None	-66.341	[9.290]***	-32.804	[6.859]***	39.089	[8.962]***	17.511	[6.817]**	32.019	[9.042]***	22.334	[5.588]***	61
Albania	-65.209	[9.231]***	-33.498	[6.918]***	38.078	[8.898]***	18.174	[6.863]***	31.566	[8.956]***	22.599	[5.595]***	60
Argentina	-66.596	[9.348]***	-32.976	[6.901]***	39.407	[9.006]***	16.986	[6.855]**	32.245	[9.111]***	22.904	[5.630]***	60
Armenia	-63.817	[9.182]***	-32.488	[6.949]***	36.660	[8.879]***	17.166	[6.905]**	30.975	[8.935]***	22.285	[5.593]***	60
Azerbaijan	-65.239	[9.245]***	-33.386	[6.925]***	38.117	[8.920]***	18.068	[6.874]***	31.587	[8.977]***	22.648	[5.597]***	60
Bangladesh	-66.873	[9.378]***	-32.768	[6.927]***	39.730	[9.000]***	17.420	[6.882]**	33.706	[8.904]***	22.299	[5.794]***	60
Belarus	-66.238	[9.325]***	-32.779	[6.877]***	38.459	[8.969]***	17.598	[6.810]***	32.193	[9.034]***	22.235	[5.584]***	60
Belize	-66.511	[9.414]***	-32.554	[6.861]***	39.503	[9.101]***	17.253	[6.834]**	31.518	[9.230]***	22.822	[5.581]***	60
Bolivia (Plurinational State of)	-66.506	[9.338]***	-32.854	[6.891]***	39.161	[9.009]***	17.775	[6.860]***	32.233	[9.087]***	22.315	[5.648]***	60
Bosnia and Herzegovina	-62.505	[9.286]***	-31.276	[7.143]***	35.778	[8.928]***	16.218	[7.048]**	30.590	[8.956]***	21.853	[5.625]***	60
Brazil	-66.097	[9.273]***	-32.613	[6.832]***	39.398	[8.945]***	17.997	[6.797]***	31.497	[9.000]***	21.751	[5.574]***	60
Bulgaria	-66.896	[9.527]***	-33.134	[7.105]***	39.678	[9.189]***	17.860	[7.033]**	32.388	[9.099]***	22.437	[5.651]***	60
Cambodia	-59.600	[8.966]***	-32.151	[6.724]***	32.481	[8.656]***	16.605	[6.692]**	28.977	[8.672]***	22.002	[5.484]***	60
Canada	-65.690	[9.385]***	-32.674	[6.929]***	38.300	[9.112]***	17.386	[6.914]**	30.046	[9.945]***	22.033	[6.058]***	60
Chile	-66.101	[9.328]***	-32.799	[6.885]***	38.323	[8.982]***	16.797	[6.842]**	32.414	[9.056]***	22.905	[5.615]***	60
China	-68.120	[9.821]***	-33.284	[7.189]***	40.713	[9.425]***	17.959	[7.103]**	32.763	[9.204]***	22.470	[5.669]***	60
Colombia	-66.467	[9.357]***	-32.983	[6.909]***	39.208	[9.016]***	17.251	[6.879]**	32.119	[9.106]***	22.769	[5.637]***	60
Costa Rica	-66.193	[9.329]***	-32.892	[6.895]***	39.227	[8.988]***	16.933	[6.850]**	31.875	[9.083]***	22.835	[5.620]***	60
Croatia	-70.658	[9.725]***	-34.550	[7.154]***	43.113	[9.369]***	19.164	[7.063]***	33.732	[9.240]***	22.971	[5.659]***	60
Czech Republic	-66.101	[9.400]***	-35.390	[6.934]***	38.760	[9.068]***	19.817	[6.879]***	31.665	[9.023]***	23.303	[5.638]***	60
Dominican Republic	-66.207	[9.315]***	-32.729	[6.874]***	38.380	[8.979]***	16.871	[6.838]**	32.742	[9.044]***	22.844	[5.615]***	60
Ecuador	-66.213	[9.332]***	-32.681	[6.886]***	38.788	[9.007]***	17.822	[6.857]***	32.188	[9.074]***	22.042	[5.641]***	60

Table 6 (continued)

Country	Small firms		Medium-sized firms		Private or public registry				Common law (country)				Countries
					Small firms		Medium-sized firms		Small firms		Medium-sized firms		
	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	
El Salvador	-66.419	[9.294]***	-32.900	[6.869]***	38.901	[8.974]***	16.832	[6.824]**	32.271	[9.068]***	22.898	[5.610]***	60
France	-66.595	[9.316]***	-32.787	[6.874]***	38.032	[8.974]***	17.355	[6.822]**	33.068	[9.016]***	22.529	[5.588]***	60
Georgia	-66.321	[9.379]***	-34.121	[6.936]***	38.920	[9.098]***	18.631	[6.929]***	31.935	[9.046]***	22.830	[5.633]***	60
Germany	-67.343	[9.390]***	-33.178	[6.935]***	38.404	[9.064]***	17.431	[6.899]**	33.484	[9.068]***	22.819	[5.643]***	60
Guatemala	-65.682	[9.271]***	-32.476	[6.849]***	38.910	[8.948]***	17.539	[6.823]**	31.824	[9.044]***	22.115	[5.627]***	60
Haiti	-66.342	[9.336]***	-32.804	[6.889]***	39.459	[9.014]***	17.778	[6.862]***	32.000	[9.104]***	22.362	[5.638]***	60
Honduras	-66.101	[9.275]***	-32.644	[6.849]***	39.121	[8.953]***	17.534	[6.821]**	31.781	[9.064]***	22.164	[5.624]***	60
Hungary	-66.165	[9.247]***	-32.789	[6.833]***	39.706	[8.947]***	17.113	[6.777]**	31.266	[9.041]***	22.417	[5.583]***	60
India	-69.311	[9.794]***	-34.094	[7.160]***	42.014	[9.563]***	18.714	[7.203]***	28.801	[9.784]***	20.436	[5.992]***	60
Indonesia	-65.691	[9.190]***	-32.386	[6.794]***	39.322	[8.884]***	18.100	[6.748]***	31.445	[9.002]***	21.752	[5.572]***	60
Italy	-66.129	[9.238]***	-32.723	[6.824]***	38.616	[8.935]***	17.291	[6.810]**	32.054	[9.023]***	22.246	[5.617]***	60
Kazakhstan	-66.096	[9.499]***	-30.476	[6.858]***	38.859	[9.145]***	15.353	[6.814]**	31.765	[9.078]***	21.224	[5.535]***	60
Kyrgyzstan	-64.962	[9.336]***	-30.766	[6.898]***	37.794	[8.999]***	15.613	[6.847]**	31.578	[9.026]***	21.578	[5.559]***	60
Lithuania	-66.483	[9.299]***	-33.115	[6.868]***	40.158	[8.963]***	16.769	[6.811]**	31.455	[9.080]***	23.094	[5.567]***	60
Malaysia	-66.361	[9.412]***	-32.919	[6.921]***	39.178	[9.127]***	17.730	[6.895]**	32.436	[9.988]***	23.181	[5.856]***	60
Mexico	-65.713	[9.253]***	-32.400	[6.836]***	39.433	[8.920]***	18.183	[6.804]***	31.453	[9.019]***	21.647	[5.606]***	60
Moldova (Republic of)	-66.820	[9.723]***	-30.081	[6.964]***	39.700	[9.347]***	15.086	[6.899]**	32.156	[9.161]***	21.176	[5.544]***	60
Nicaragua	-65.936	[9.265]***	-32.535	[6.838]***	39.364	[8.951]***	17.743	[6.807]***	31.606	[9.056]***	21.981	[5.606]***	60
Pakistan	-67.705	[9.237]***	-32.813	[6.854]***	40.812	[8.843]***	17.495	[6.812]**	37.300	[7.796]***	23.135	[5.526]***	60
Panama	-66.117	[9.300]***	-32.738	[6.866]***	39.102	[8.954]***	17.067	[6.833]**	31.732	[9.061]***	22.455	[5.616]***	60
Peru	-66.768	[9.337]***	-33.077	[6.900]***	38.755	[9.012]***	17.224	[6.871]**	32.439	[9.093]***	22.629	[5.646]***	60
Philippines	-66.590	[9.298]***	-32.828	[6.861]***	39.089	[8.961]***	17.942	[6.826]***	31.928	[9.067]***	21.941	[5.625]***	60
Poland	-68.628	[9.483]***	-33.771	[7.079]***	41.276	[9.119]***	18.447	[7.003]***	32.928	[9.118]***	22.642	[5.628]***	60

Table 6 (concluded)

Country	Small firms		Medium-sized firms		Private or public registry				Common law (country)				Countries
					Small firms		Medium-sized firms		Small firms		Medium-sized firms		
	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	Coef.	Standard deviation	
Portugal	-65.849	[9.275]***	-32.494	[6.847]***	38.539	[8.928]***	17.569	[6.817]***	32.084	[9.028]***	22.020	[5.585]***	60
Romania	-65.863	[9.537]***	-35.259	[7.111]***	38.767	[9.150]***	19.875	[6.998]***	31.825	[9.021]***	23.375	[5.664]***	60
Russian Federation	-68.005	[9.238]***	-31.425	[6.672]***	40.366	[8.920]***	16.070	[6.622]**	32.710	[8.997]***	21.747	[5.431]***	60
Singapore	-68.972	[9.670]***	-34.366	[7.101]***	41.971	[9.372]***	19.288	[7.109]***	28.040	[9.350]***	19.775	[5.649]***	60
Slovakia	-65.449	[9.233]***	-32.488	[6.829]***	40.305	[8.945]***	17.653	[6.790]***	30.502	[9.039]***	21.951	[5.570]***	60
Slovenia	-66.099	[9.358]***	-32.674	[6.897]***	38.663	[9.026]***	17.986	[6.865]***	31.930	[9.053]***	21.925	[5.631]***	60
Spain	-65.977	[9.283]***	-32.430	[6.842]***	38.902	[8.944]***	18.208	[6.794]***	31.960	[9.042]***	21.594	[5.609]***	60
Sweden	-66.638	[9.294]***	-32.855	[6.861]***	37.854	[8.912]***	16.964	[6.775]**	33.197	[8.971]***	22.835	[5.570]***	60
Thailand	-67.840	[9.696]***	-33.078	[7.152]***	39.766	[9.372]***	17.119	[7.192]**	33.938	[9.389]***	24.573	[5.803]***	60
Trinidad and Tobago	-67.552	[9.550]***	-33.831	[7.014]***	40.075	[9.273]***	18.598	[6.988]***	33.457	[10.267]***	24.998	[5.969]***	60
Turkey	-66.361	[9.337]***	-32.777	[6.878]***	38.766	[9.016]***	17.510	[6.851]**	32.243	[9.040]***	22.295	[5.609]***	60
United Kingdom	-66.474	[9.349]***	-32.508	[6.844]***	39.454	[9.029]***	17.210	[6.812]**	33.038	[9.360]***	21.487	[5.563]***	60
United States	-65.945	[9.440]***	-32.599	[6.923]***	38.786	[9.151]***	17.310	[6.902]**	31.529	[9.914]***	21.766	[5.903]***	60
Ukraine	-65.419	[9.432]***	-31.247	[6.947]***	38.163	[9.077]***	16.021	[6.878]**	31.608	[9.003]***	21.692	[5.532]***	60
Uruguay	-66.776	[9.322]***	-32.935	[6.883]***	38.825	[8.988]***	17.867	[6.851]***	32.412	[9.074]***	22.184	[5.646]***	60
Uzbekistan	-68.264	[9.564]***	-32.801	[6.990]***	40.889	[9.204]***	17.536	[6.913]**	32.756	[9.143]***	22.263	[5.599]***	60
Venezuela (Bolivarian Republic of)	-66.196	[9.288]***	-32.778	[6.860]***	39.194	[8.965]***	17.683	[6.817]***	31.915	[9.054]***	22.286	[5.611]***	60

Source: World Bank, World Business Environment Survey (WBES) and S. Djankov, C. McLiesh and A. Shleifer, "Private credit in 129 countries", *Journal of Financial Economics*, vol. 84, No. 2, Amsterdam, Elsevier, 2007.

Note: Robust standard errors shown in brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

VII. Final remarks

Using a sample of 61 countries, this study provides empirical evidence that points to the importance of creditor protection and the development of information-sharing mechanisms in opening up access to bank credit for firms of different sizes. We use limited dependent variable techniques to test if the development of information-sharing institutions, such as private credit bureaux or public credit registries, helps to reduce the financing gap between large companies and small and medium-sized enterprises. Using country dummies, we control for any country-omitted variable, which has been one of the major problems in previous studies on this topic.

Our results suggest that improving the coverage of private credit bureaux and public credit registries has a statistically significant effect in terms of a reduction in the gap between the percentage of investment financed with bank credit for large firms and small firms. The results obtained in this study are not only statistically significant, but are also economically meaningful. A one-standard-deviation increase in the coverage of credit bureaux reduces the financing gap between small firms and large ones by eight percentage points. Using the same technique, we test if the protection of creditors, measured by effective creditor rights and the type of legal system, reduces the gap between bank credit for large firms and small firms. In common-law countries, which are characterized by a high level of creditor protection, this gap is between 20 and 30 percentage points less than it is in other countries. An improvement in effective creditor rights of one standard deviation reduces the gap between small and large firms by from 9 to 14 percentage points.

The main policy driver for the reduction of the financing gap is an improvement in the protection of creditor rights. Especially for small firms, however, efforts to strengthen creditor rights, which constitutes a titanic task entailing widespread reforms, can be bolstered by the development of information-sharing mechanisms. Our results suggest that the pay-off of this latter type of effort will be greater in countries where creditors' rights are poorly protected. Future theoretical and empirical works would do well to study the complementarity of information sharing and creditor property rights as reflected in our results. Another interesting avenue of research would be to focus on the effect of information sharing on competition and vice versa and on the impact of this interaction on access to credit for SMEs.[23] As shown in this study, information sharing reduces information asymmetries between lenders and creditors, but this may also reduce banks' ability to appropriate SME rents. This may increase these firms' profits but, as claimed by Petersen and Rajan (1994), it may also make banks less willing to lend to SMEs and therefore reduce their access to credit.

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Annex A1

Table A1.1
Country-level data

	Average share of investment financed with bank credit	Public or private credit information institution dummy	Public credit registry coverage (per 1,000)	Private credit bureau coverage (per 1,000)	Common law regime	Effective creditor rights index	GDP per capita, PPP adjusted (*in logs*)
Albania	3	0	0	0	0	-0.62	8.39
Argentina	30	1	149	475	0	-0.01	9.25
Armenia	3	0	0	0	0	-0.25	7.69
Azerbaijan	2	0	0	0	0	-0.73	7.74
Bangladesh	13	1	1	0	1	-0.39	6.76
Belarus	5	1	0	0	0	-0.59	8.61
Belize	32	0	0	0	1	0.16	8.56
Bolivia (Plurinational State of)	24	1	55	134	0	-0.3	8.13
Bosnia and Herzegovina	16	0	0	0	0	-0.61	8.47
Brazil	26	1	44	439	0	-0.06	8.97
Bulgaria	6	0	0	0	0	-0.03	8.78
Cambodia	7	0	0	0	0	-0.46	6.85
Canada	21	1	0	806	1	0.53	10.34
Chile	38	1	209	227	0	0.71	9.25
China	9	0	0	0	0	-0.18	7.82
Colombia	29	1	0	187	0	0	8.79
Costa Rica	18	1	7	55	0	0.21	8.97
Croatia	19	0	0	0	0	-0.07	9.39
Czech Republic	10	0	0	0	0	0.53	9.71
Dominican Republic	26	1	0	423	0	-0.17	8.6
Ecuador	15	1	82	0	0	0	8.64
El Salvador	28	1	130	128	0	-0.32	8.53
France	11	1	12	0	0	0	10.25
Georgia	7	0	0	0	0	-0.47	7.73
Germany	17	1	5	693	0	1.55	10.3
Guatemala	27	1	0	35	0	-0.21	8.27
Haiti	11	1	1	0	0	-0.77	7.08
Honduras	20	1	45	0	0	-0.43	7.97
Hungary	15	1	0	15	0	0.22	9.47
India	33	0	0	0	1	0.06	7.41
Indonesia	15	1	3	0	0	-0.44	7.92
Italy	42	1	55	416	0	0.51	10.2
Kazakhstan	7	0	0	0	0	-0.67	8.55
Kyrgyzstan	1	0	0	0	0	-0.63	7.29
Lithuania	8	1	7	0	0	0.1	9.14
Malaysia	17	1	105	461	1	0.57	9.2
Mexico	11	1	0	382	0	0	9.35
Moldova (Republic of)	7	0	0	0	0	-0.2	7.31

Table A1.1 (concluded)

	Average share of investment financed with bank credit	Public or private credit information institution dummy	Public credit registry coverage (per 1,000)	Private credit bureau coverage (per 1,000)	Common law regime	Effective creditor rights index	GDP per capita, PPP adjusted (in logs)
Nicaragua	17	1	50	0	0	-0.86	7.62
Pakistan	27	1	1	0	1	-0.17	7.55
Panama	44	1	0	302	0	0.04	8.98
Peru	25	1	92	185	0	0	8.61
Philippines	19	1	0	22	0	-0.08	7.84
Poland	13	0	0	0	0	0.16	9.32
Portugal	13	1	496	24	0	0.35	9.89
Romania	10	0	0	0	0	-0.13	8.85
Russian Federation	6	0	0	0	0	-0.23	9
Singapore	23	0	0	0	1	1.71	10.46
Slovakia	11	1	2	0	0	0.13	9.44
Slovenia	17	1	14	0	0	0.7	9.84
Spain	20	1	305	48	0	0.71	10.09
Sweden	19	1	0	489	0	0.55	10.23
Thailand	34	0	0	0	1	0.21	8.56
Trinidad and Tobago	37	1	315	0	1	0.11	9.42
Turkey	20	1	7	0	0	0.03	9.12
United Kingdom	11	1	0	652	1	2.14	10.25
United States	18	1	0	810	1	0.49	10.54
Ukraine	6	0	0	0	0	-0.42	8.2
Uruguay	32	1	49	479	0	0.32	9.14
Uzbekistan	5	0	0	0	0	-0.59	7.37
Venezuela (Bolivarian Republic of)	15	1	97	0	0	-0.67	9.19

Source: Prepared by the authors, on the basis on World Bank, World Business Environment Survey (WBES) and S. Djankov, C. McLiesh and A. Shleifer, "Private credit in 129 countries", *Journal of Financial Economics*, vol. 84, No. 2, Amsterdam, Elsevier, 2007.

Mechanisms of default risk transmission and economic policy coordination

Karlo Marques Junior and Fernando Motta Correia

Abstract

This paper analyses the coordination between monetary and fiscal policy in an emerging economy with an inflation-targeting monetary regime, in a context in which default risk shocks can lead to macroeconomic imbalances. It develops a macrodynamic model in order to capture the mechanisms of default risk transmission and its effects on the definition of reaction functions for the monetary and fiscal authorities. The main findings of the model point to the existence of new mechanisms of default risk transmission associated with price and fiscal stability.

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I.   Introduction

Given the consensus that low, stable inflation is desirable (Carlin and Soskice, 2006), many countries opt for inflation-targeting regimes. Their experiences have shown that the operating framework of economic policy includes not only measures of a monetary nature, but also the quest for equilibrium in the public accounts.

Among the authors who draw attention to the relationship between fiscal and monetary policy are Sargent and Wallace (1981), Woodford (1995, 1996 and 2001), Leeper (1991 and 2009) and a number who examine the case of Brazil, such as Favero and Giavazzi (2003) and Blanchard (2004).

The main question is whether, on the one hand, the adverse behaviour of some fiscal variables may prevent monetary policy from functioning properly under an inflation-targeting regime with a monetary policy rule similar to Taylor's (1993). On the other hand, the functioning of monetary policy also affects the performance of fiscal policy.

In an emerging economy, risk perception can affect and be affected by both monetary and fiscal policy. The risk attributed to each country depends largely on its sovereign bond solvency. When a rising interest rate pushes up the cost of debt servicing, default risk may be expected —all else remaining constant— to increase as well. This creates a cyclical effect, since financial investors will accordingly demand a higher premium to buy sovereign bonds.

It is important to look at some aspects of the effect that exogenous economic shocks —such as risk shocks— have on public debt, because inflation-targeting monetary regimes use the nominal interest rate to control the shocks an economy is exposed to. So the effect of monetary policy on public debt is passed through variables such as public debt composition and crises of confidence.

Initially, the effects of monetary policy on debt are associated with shocks that can push inflation off target and thus induce the monetary authority to raise interest rates to cancel out the effects of the shock. Since the monetary authority is free to change the interest rate, debt rises as a function of inflationary volatility. In addition, confidence shocks can occur at times of great uncertainty, when investors tend to flee from higher risk or demand higher yields in compensation, which again pushes up the cost of debt servicing.

In both cases, then, monetary policy affects debt, which can subsequently push up default risk even more and thus trigger the vicious cycle mentioned above. This is why a better understanding is needed of the relationship between the conduct of monetary policy and the behaviour of public debt in settings where sovereign risk can lead to macroeconomic disequilibria.

By the same token, following Sargent and Wallace (1981), if the fiscal authority does not respect intertemporal budgetary constraints, the monetary authority can find itself in difficulty, having to finance debt servicing through inflation tax, thus losing control over inflation.

This article analyses how the coordination of monetary and fiscal policies should be managed in an emerging economy with an inflation-targeting regime, in a context in which default risk shocks can lead to macroeconomic disequilibria. The main contributions of this work are: (i) suggesting a monetary policy rule that includes a fiscal variable, specifically the deviation of public debt from a target, and (ii) establishing a model in which risk is a factor that can lead to macroeconomic disequilibria[1] through its effects on the exchange rate, on the interest rate and, indirectly, on public debt and inflation. In other words, the article aims to answer the following question: what considerations should the monetary authority bear in mind regarding its policies' impact on the variables targeted by the fiscal authority, and vice versa, in an economic setting such as that described?

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[1]   In this work, macroeconomic disequilibrium is defined as a situation in which the interest rate, the debt-to-GDP ratio and default risk do not converge over time towards a stable equilibrium.

The article seeks to understand how macroeconomic policies should adapt to exogenous shocks in order to maintain stable equilibrium in different economic circumstances. This involves developing a monetary policy model that could be adopted to minimize the impact of shocks on debt risk, both when the interest rate is used to guide inflation towards set targets, and when fiscal policy targets a particular balance on the public accounts.

This article has six sections. After this Introduction, the second section reviews the theory on which the models used are based. The third section sets out the hypothesis and the comparative statics analysis of the model, in order to analyse the short-term effects of risk on important economic policy variables. The fourth section studies the model's dynamic stability in a configuration consisting of a budgetary rule set by the fiscal authority in coordination with the monetary authority. In the fifth section, the model is extended by introducing risk perception as an explanatory variable for the exchange rate, with a view to capturing the effects of risk shocks on macroeconomic equilibrium more precisely. Here, default risk is part of the function that determines —and can thus have undesired effect on— the exchange rate. Lastly, the sixth section sets forth the article's conclusions, which summarize the theoretical corollaries suggested in the model.

II. Coordination between monetary and fiscal policies: a literature review

A number of works draw attention to the link between fiscal policy and monetary policy and indicate that the central bank cannot ignore the influence of fiscal variables on price levels in an inflation-targeting regime that seeks a monetary reaction in line with Taylor (1993). Since, moreover, the functioning of monetary policy affects fiscal policy, coordination between the two is necessary. In setting the interest rate, then, the central bank must keep in mind the trajectory of debt over time.

According to Blanchard, Dell'Ariccia and Mauro (2010), in the 1960s and 1970s monetary and fiscal policies were equally important and were generally treated as two instruments for achieving the twofold objective of domestic and external equilibrium. In the intervening decades, however, fiscal policy tended to take a back seat as policymakers' attention turned mainly to monetary matters. Nevertheless, several authors signalled the limitations of monetary policy, especially in hostile fiscal conditions. Some of these works are mentioned below.

Sargent and Wallace (1981) state that monetary policy could become ineffective to control inflation if the fiscal authority disregards the government's long-term intertemporal budget constraint. This would occur because of the need to finance public deficits through seignorage, which generates inflation tax. Depending on the interaction of economic policies, the authors distinguish two situations: monetary dominance and fiscal dominance.

Monetary dominance obtains when the fiscal authorities' spending is limited by the bond demand function, such that a fiscal surplus is needed to keep a constant ratio between public sector debt and GDP. In this situation, the monetary authorities determine the money supply and government spending is constrained by that decision. With this type of policy coordination, the monetary authorities control inflation through the supply of base money, thus setting up a scenario of monetary dominance in which monetary policy is said to be active and fiscal policy, passive.

In conditions of fiscal dominance, the fiscal authority sets its budgets independently, announcing all current and future deficits and surpluses and thus determining the amount of revenue that must be raised through bond sales and seignorage (Sargent and Wallace, 1981). Here, the fiscal authorities take no account of the need for a large enough surplus to keep the debt-to-GDP ratio under control.

This second type of coordination produces what Sargent and Wallace term "unpleasant arithmetic." The monetary authorities become passive and lose control over inflation, because they are forced to increase seignorage revenue to keep the government solvent. This is particularly prejudicial when investors' demand for bonds, balanced with the government's demand for liquid resources, entails an interest rate higher than the economic growth rate. Nevertheless, although the inflation is generated by a fiscal imbalance, it remains a monetary phenomenon.

The work of Leeper (1991) is another important reference in the literature on coordination between monetary and fiscal policies. Leeper defines four situations in which monetary policy may be considered active or passive, depending on its responsiveness to public debt shocks.

In Leeper's model (1991), policy authorities can set a control variable actively, while an intertemporally balanced government budget requires that at least one authority sets its control variable passively. When both policies are passive, policy is incompletely specified and the price function is indeterminate. If both policies are active, independent variations are allowed, violating the government budget constraint.

The notion of inflation as a purely monetary phenomenon was later questioned by economists whose ideas became formally grouped under what is known as the fiscal theory of the price level, among them Cochrane (1998 and 2001) and Woodford (1994, 1995 and 2001). According to Kocherlakota and Phelan (1999), under this theory, the definition of price growth as simply the difference over time between growth in the money supply and growth in output is severely flawed. The point they make is that the amount of money agents wish to keep in the present depends fundamentally on their expectations of inflation in the future. This opens up possibility of multiple equilibria in the trajectory of inflation, beyond the relationship between the money supply and the quantity of goods produced. Accordingly, Taylor-type monetary policy rules (1993) alone are not enough to control price levels.

According to Basseto (2008), "the fiscal theory of the price level (FTPL) describes fiscal and monetary policy rules such that the price level is determined by government debt and fiscal policy alone, with monetary policy playing at best an indirect role." In that case, prices are determined by the fiscal authority through the government budget constraint. As part of this theory, Basseto (2008) also argues that the role of the monetary authority in pricing becomes evident when the interest rate affects the evolution of nominal public debt.

A number of works on fiscal dominance focus on the role of default risk as a mechanism by which the monetary authority of an emerging country running an inflation-targeting regime can lose control of price levels. Two major works in this line refer to the Brazilian economy in the period around the 2002 elections: Favero and Giavazzi (2003) and Blanchard (2004).

Favero and Giavazzi (2003) emphasize Brazil's highly volatile country risk between 2002 and 2003, and the way in which some economic variables, especially the exchange rate, fluctuated in parallel with risk. This pointed to a vicious cycle propagated in the Brazilian economy by default risk. The authors found that a rise in default risk led —because of the interruption in capital flows— to exchange-rate devaluation and a rise in the debt-to-GDP ratio, which was heavily indexed to the dollar at that time. Devaluation and rising debt pushed up expectations of inflation and thus the monetary policy interest rate as well.

Blanchard (2004) examines the effects of tight monetary policy in an inflation-targeting regime in the context of a high debt-to-GDP ratio, strong indexation of the public debt to foreign currencies and a high degree of risk aversion on the part of international investors.

In these circumstances, an inflation shock leads to a rise in the interest rate and thus in the ratio between net government debt and GDP. This sequence will likely increase risk perception, triggering a capital flight that produces exchange-rate devaluation and, ultimately upward pressure on inflation.

Thus the situation is one of fiscal dominance similar to that described by Favero and Giavazzi (2003), whereby monetary policy is ineffective in controlling inflation in a situation of fiscal imbalance amid high risk aversion. Emerging economies under an inflation-targeting regime may be particularly prone to finding themselves in this perverse situation, insofar as investors consider their bonds a high-risk portfolio choice.

In this scenario, monetary policy loses control over inflation and becomes dominated by expectations regarding fiscal conditions. Blanchard (2004) suggests empirically that the Brazilian economy exhibited such fiscal dominance between 1999 and 2004 and that default risk was the factor that triggered these economic policy interactions.

It may be deduced from these reflections on the combination of macroeconomic policies that the Taylor rule has limitations in terms of ensuring long-term economic stability, especially in emerging economies.

It is important to establish the nature of the default risk transmission mechanism in emerging economies, particularly given their exposure to exchange-rate shocks. Considering this mechanism, the original Taylor rule seems limited in its ability to guide the action of the monetary authority.

III. Risk, interest and debt: identifying default risk transmission mechanisms

Following the discussion above with respect to macroeconomic policy coordination, a macroeconomic model is presented below that seeks to represent hypotheses for an emerging economy with an inflation-targeting regime, in which economic policy coordination is aimed at minimizing the adverse effects of risk shocks.

A few additional considerations are called for before presenting the basic structure of the model. Since economies that rely on inflation-targeting regimes may be exposed to exchange-rate shocks, insofar as exchange-rate flexibility is a basic condition for such a monetary regime, many authors include the exchange rate in their analysis as a variable in the reaction function of the central bank. Under the assumption of an open economy with an inflation-targeting monetary regime, Ball (1999) suggests that optimal monetary policy should incorporate a monetary conditions index comprising the interest rate, the exchange rate and a measure for the inflation target.

However, Ball's analysis (1999) has some limitations when it comes to emerging economies, where —as noted earlier— high risk perceptions can compromise the objective of monetary policy. In such economies, the risk rating is positively correlated with the exchange rate.[2] So, in emerging economies such as Brazil, a reaction function such as the Taylor rule could increase risk perception through interest rate rises, and lead to frequent exchange-rate shocks, as suggested by Favero and Giavazzi (2003) and Blanchard (2004).

Accordingly, in the case of an economy in which the fragility of certain fiscal variables could affect the conduct of monetary policy because of the link between the risk premium and the exchange rate, there are grounds for suggesting that central banks react to agents' risk perception.

The magnitude of the risk premium thus incorporates the uncertainties built into the commitment to remunerate the public debt until its maturity. It must always be recalled that the shocks to which

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[2]   Svensson (2000) and Ball (1999) assume that in industrialized countries the risk premium follows a random walk that does not affect the conduct of monetary policy, whereas in developing countries there is a strong link between risk perception and capital flight, which has effects on the exchange rate and on inflation. This argument complements the definition of emerging economy given earlier.

the structure is exposed through the maturity structure of the interest rate make the risk premium component subject to expectations shocks, bearing in mind that, where fiscal policy is not committed to a stable debt-to-GDP ratio, agents can demand a high rate of return to compensate for the high risk of assuming a debt with great default probability.

In this context, this work develops a model with three non-homogenous, simultaneous first-order linear differential equations, in order to study the coordination between fiscal and monetary policies. Its long-term variables are the risk premium, the nominal interest rate and the behaviour of the debt-to-GDP ratio. Thus, an analysis is made of the long-run behaviour of the variables, that is, whether or not they converge over time with their steady state, i.e. if the equilibrium is dynamically stable. It should be recalled that long-run risk stability is the desired outcome in order to guarantee macroeconomic stability and monetary policy effectiveness in an inflation-targeting regime.

1.   Model of default risk transmission mechanisms

According to the maturity structure of the interest rate, the rate of return on a bond at time t depends on the average short-term interest rate during its term n, plus a risk premium given by the market conditions for that bond. Thus, the relationship between short- and long-term interest rates may be rendered as follows:

$$r_{nt} = \frac{r_t + r^e_{t+1} + r^e_{t+2} + r^e_{t+3} + \ldots + r^e_{t+(n-1)}}{n} + R_{nt}$$

where r_{nt} denotes the bond's real long-term interest rate at maturity, r_t the real short-term interest rate for period t and r^e_t the real interest rate expected for period t.

The maturity structure of the interest on the public debt security can be simplified as follows:

$$r = r^e + R \tag{1}$$

Equation (1) decomposes the return on government bonds into two components: first, the expectations of short-term real interest rates up to maturity (r^e) and, second, R, the risk premium to which bond buyers are exposed.

In equation (1), R is a measure of default risk that captures the uncertainties relating to the commitment to pay the return on the bond until its maturity. In general, the longer the maturity, the higher both yield and risk. Agents' perceptions of the magnitude of default risk variation depend on the comparison of a bond that pays rate r with respect to a risk-free security, in this case denoted $\bar{i}$. Since $\bar{i}$ represents the nominal rate on a risk-free security,[3] it may be assumed that the variation in risk R over time is reflected in the difference between those two rates, that is, the difference between rate r and rate $\bar{i}$. The idea is that this gap arises from the risk compensation demanded by agents, so that, over the long term, the larger this difference, the larger the intertemporal variation in default risk, as shown in the following differential equation (2):

$$\dot{R} = \sigma(r - \bar{i}) \quad , \sigma > 0 \tag{2}$$

Thus, coefficient σ captures the sensitivity of default risk R to the spread between rates of return on positive-risk bonds and risk-free bonds. Analogously, σ captures economic agents' risk aversion. This is expected to show a direct relation to the debt-to-GDP ratio.

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[3]   In general, United States Treasury bonds are considered risk-free bonds for international investors.

The nominal interest rate in the economy in question is defined by real interest rate (r) plus the inflation rate (π), as in the relation represented by equation (3) below, similar to Fisher's rule:

$$i = r + \pi \tag{3}$$

This equation suggests that the nominal interest rate can vary as a result of changes in both the real interest rate and in the inflation rate.

It is assumed that the short-term nominal interest rate is set by the central bank (i^*), as the main monetary policy tool for guiding inflation towards the desired target, and that it is differentiated from i as being the base rate desired by the monetary authority, that is:

$$i = i^* \tag{4}$$

Non-Ricardian public debt behaviour can also cause the monetary authority to lose control over inflation. Accordingly, the success of an inflation-targeting regime may also require the fiscal authority to set a target for the proportion of net debt over GDP to ensure debt solvency in the long run. The central bank's reaction function can thus take into account fiscal shocks in the economy.

Three factors thus influence the central bank's decision to set the nominal interest rate intertemporally in an emerging economy: on the one hand, when inflation (π) diverges from the preset target (π^*), the monetary authority reacts to contain that deviation. On the other hand, since that rate also determines the return on government bonds, as in equation (3), it is assumed that the interest rate must react to deviations in public debt (b) from a target (b^*) set according to economic policy guidelines aimed at maintaining public debt sustainability, that is, to equalize government expenditures and revenues at present values.[4] In that sense, a component of present policy coordination is presupposed in the reaction function of the monetary authority.

This nominal interest rate reaction occurs because the prospect of debt insolvency would force the monetary authority to resort to inflation tax, thus losing control of inflation. It also occurs because public debt produces an autonomous effect on the risk premium (R) when it diverges from a preset target, acting as a kind of thermometer for investors as to the risk of default on government bonds. It will be recalled that a rise in default risk can trigger capital flight and thus exchange-rate depreciation, which in turn produces inflationary pressure.

The third component of the monetary authority's reaction function is the spread between the nominal interest rate in the domestic market (i^*) and the nominal interest rate in the external market ($\bar{i}$). The larger this spread, the less the need for the monetary authority to raise its own domestic interest rate, bearing in mind that, if domestic interest rates remain constant, reduction in the external interest rate stimulates exchange-rate appreciation, which aids price stability. The idea, then, is to capture the indirect effect of the exchange rate on inflation, since the interest rate spread should determine that rate. The parameter μ below thus refers to the central bank's concern with exchange-rate variations. This dynamic is set forth in the following differential equation, which is an adaptation of Taylor's rule (1993):

$$\frac{di}{dt} = \beta\left(\pi - \pi^*\right) + \alpha\left(b - b^*\right) + \mu\left(i^* - \bar{i}\right) \ , \beta > 0; \alpha > 0; \mu < 0 \tag{5}$$

The suggested monetary rule described above is one of the main contributions of this work. If it is accepted that Taylor's rule is not the best suited to an emerging economy, a fiscal variable must be

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[4] The fiscal authority could stipulate target b^* in coordination with the monetary authority to determine macroeconomic policies to reach the inflation target.

included in the monetary policy rule adopted by the central bank. Section V makes another important contribution, by identifying the channel through which default risk can destabilize the model presented.

In light of the above, it is necessary to establish the difference between the last two equations, given that equation (4) reflects the exogenous short-term nature of the interest rate in an inflation-targeting regime. Unlike equation (5), which specifies a monetary policy rule that will limit the monetary authority's decisions on setting the base rate of interest over a given period, some fluctuations in the base rate can occur over short periods, as defined in (4).

Equation (6) shows the government's intertemporal constraint:

$$\dot{b} = ib + g - t \tag{6}$$

where g is government expenditure and t its revenues. Considering a public debt indexed to the nominal interest rate, an increase in i will have an incremental effect on debt b, and produce primary deficits in the public sector ($g - t > 0$).

The inflation rate is, *a priori*, determined by an expectations-augmented Phillips curve, to which a component representing the nominal exchange rate is added:[5]

$$\pi = \tau(y - \bar{y}) + \pi^e + \theta(E) \quad, \tau > 0; \theta > 0 \tag{7}$$

where $(y - \bar{y})$ represents the output gap, π^e expected inflation and E the nominal exchange rate.[6]

In this case, the exchange-rate effect is included as well as the traditional effects on inflation represented by the Phillips curve. Empirical tests conducted by Goldfajn and Werlang (2000) and Correa and Minella (2010) found that the pass-through effect is sharper when: (i) the economy is in a rapid expansion cycle; (ii) exchange-rate volatility is low; (iii) the economy is very open; (iv) the initial rate of inflation is high, especially —according to the first work cited— in emerging economies, and (v) exchange-rate misalignment is severe.

However, the expectations component (π^e) is determined by the deviations of expected output (y^e) with respect to potential output $(\bar{y})$, and by the difference between the expected nominal exchange rate[7] and its equilibrium level $(E^e - E^*)$, since, as noted earlier, the exchange rate affects price levels, especially when it is above equilibrium. Thus:

$$\pi^e = \phi(y^e - \bar{y}) + \varkappa(E^e - E^*), \phi > 0; \varkappa > 0 \tag{8}$$

The idea behind the expectations component is that agents make their forecasts of inflation by observing the equilibrium between aggregate supply and demand. Their expectations also incorporate forecasts on the exchange-rate, which is an important variable in price composition. Under the hypothesis of rational expectations, agents forecast output behaviour expecting that, in the absence of exogenous shocks, observed output will be equal to potential output and the exchange rate will be equal to the equilibrium rate. Accordingly, assuming that $y^e = \bar{y}$ and $E = E^*$, (8) may be substituted into (7) to give the following Phillips curve:

$$\pi = \tau(y - \bar{y}) + \theta(E) \tag{7.1}$$

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[5]   The effect of the nominal exchange rate on inflation was considered of negligible magnitude in early works with the Phillips curve. However, several works have estimated it empirically.

[6]   More precisely, inflation is a dependent function of the rate of exchange-rate devaluation. Equation (9), which denotes the parity of the interest rate and the exchange rate, should also be expressed in terms of exchange rate devaluation, assuming the form $\dot{E}/E = (i - \bar{i}) + R$. However, the simplification used does not affect the model's results.

[7]   The exchange rate presented is the ratio between the real and the dollar, i.e. national currency/foreign currency.

In determining the real exchange rate, e is assumed, for simplicity's sake, to be the equilibrium between domestic and external prices $\left(p = p^* \right)$. Consequently, the exchange rate is determined by interest rate parity, according to equation (9):

$$E = E^* = e = \rho\left(i^* - \bar{i}\right) \quad \rho < 0 \tag{9}$$

In turn, aggregate demand is composed of the consumption function, the investment function, government spending and the trade balance, as expressed in equation (10), which denotes an IS curve for an open economy:

$$y = c\left(y\right) + I\left(i\right) + g + x\left(E\right), c_y > 0, I_i < 0, x_e > 0 \tag{10}$$

2.   Short-term analysis: a comparative statics study

A comparative statics analysis is now presented in order to deduce some short-term relations between the model's key variables.

By substituting (4) into (10), we obtain:

$$y = \left(\frac{I_i}{1 - c_y}\right)i^* + \left(\frac{1}{1 - c_y}\right)g + \left(\frac{X_e}{1 - c_y}\right)E \tag{10.1}$$

As inflation is a function of income, equation (10.1) is substituted into (7.1), giving the following behaviour for inflation:

$$\pi = \left(\frac{\tau I_i}{1 - c_y}\right)i^* + \left(\frac{\tau}{1 - c_y}\right)g + \left(\frac{\tau X_e}{1 - c_y} + \theta\right)E - \tau\bar{Y} \tag{7.2}$$

Inserting equations (1) and (4) into (3), and the result into (7.2), then inserting (9) in the final result gives:

$$\pi = \left[\frac{\left(\frac{\tau I_i}{1 - c_y}\right)}{v} + \frac{\left(\theta + \frac{\tau X_e}{1 - c_y}\right)}{v}(\rho)\right]R + \left[\frac{\left(\frac{\tau I_i}{1 - c_y}\right)}{v} + \frac{\left(\theta + \frac{\tau X_e}{1 - c_y}\right)}{v}\rho\right]r^e$$

$$- \frac{\left(\theta + \frac{\tau X_e}{1 - c_y}\right)}{v}\bar{i}\rho + \frac{\left(\frac{\tau}{1 - c_y}\right)}{v}g - \frac{\tau\bar{Y}}{v} \tag{7.3}$$

where $v = \left[1 - \left(\frac{\tau I_i}{1 - c_v}\right) + \left(\theta + \frac{\tau X_e}{1 - c_v}\right)\rho\right] > 0$

The derivative that indicates the short-term effect of risk on inflation can be obtained from equation (7.3):

$$\frac{\partial\pi}{\partial R} = \frac{\left(\frac{\tau I_i}{1 - c_y}\right)}{v} + \frac{\left(\theta + \frac{\tau X_e}{1 - c_y}\right)}{v}(\rho) < 0 \tag{7.3.1}$$

It is concluded from this derivative that inflation and risk are negatively correlated in the short run. The risk variable is one of the components of the short-term interest rate, as shown in (1). Thus, rising default risk exerts increasing pressure on the interest rate, thereby compressing aggregate demand and, thus, inflation. Importantly, however, although this relationship is observed in the short run, risk can introduce macroeconomic instability in the long term, leading the monetary authority to lose control over the inflation target.

Diagram 1 illustrates the channel of risk pass-through to inflation. When risk rises, the interest rate on government bonds must also rise to sustain demand by international investors. Consequently, aggregate demand falls and, ultimately, so does inflation.

Diagram 1
Flow diagram 1: channel of risk pass-through to inflation
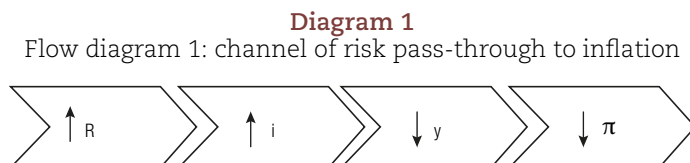
Source: Prepared by the authors.

In order to study the effect of risk changes on public debt, a function must be defined for that variable. Thus, (3) is inserted into (6), such that:

$$g - t = (r + \pi)b \tag{6.1}$$

When (1) is then substituted into (6.1), we obtain:

$$g - t = (r^e + R + \pi)b \tag{6.2}$$

Recalling that the behaviour of inflation was shown in (7.3), substituting this into (6.2) and isolating b, gives the following behaviour for public debt:

$$b = \frac{g - t}{\left\{ \left[1 + \frac{\left(\frac{\tau I_i}{1 - c_y} \right)}{v} + \frac{\left(\theta + \frac{\tau X_e}{1 - c_y} \right)}{v} \rho \right] R + \left[1 + \frac{\left(\frac{\tau I_i}{1 - c_y} \right)}{v} + \frac{\left(\theta + \frac{\tau X_e}{1 - c_y} \right)}{v} \rho \right] r^e + \frac{\left(\frac{\tau}{1 - c_y} \right)}{v} g - \frac{\left(\theta + \frac{\tau X_e}{1 - c_y} \right)}{v} i\rho - \frac{\tau}{v} \bar{Y} \right\}} \tag{6.3}$$

The effect of a marginal increase in risk on the behaviour of public debt can be analysed on the basis of the comparative statics outcomes. Below is the derivative of public debt with respect to risk, followed by some algebraic manipulations.

$$\frac{\partial b}{\partial R} = \frac{(t - g)\left[1 + \frac{\left(\frac{\tau I_i}{1 - c_y} \right)}{v} + \frac{\left(\theta + \frac{\tau X_e}{1 - c_y} \right)}{v} (\rho) \right]}{i^2} > 0 \tag{6.3.1}$$

> 0 for $t > g$ or, < 0 for $t < g$.

By algebraic reorganization and treating the marginal propensity to save as , the above derivative may be rewritten as:

$$\frac{\partial b}{\partial R} = \frac{(t-g)\left[1 + \frac{\tau I_i}{(s-\tau I_i)} + \frac{\theta + \tau X_e}{(s-\tau I_i)}(\rho)\right]}{i^2}$$

(6.3.2)

Diagram 2 shows the mechanism of risk pass-through to public debt in the event of a primary surplus. An increase in risk generates a rise in the interest rate (owing to investor's requirements) and thus increases the public debt.

<p style="text-align:center">Diagram 2
Flow diagram 2: channel of risk pass-through to public debt</p>

Source: Prepared by the authors.

The impact of the risk variable on public debt will depend on the primary balance in the government accounts. In the event of a surplus —whereby government revenues exceed expenditure— risk will have a positive effect on public debt. In the opposite case, i.e. a primary deficit, the risk impact on public debt will be negative. In addition to these considerations, the higher the marginal propensity to save and the higher the interest rate, the smaller the impact of risk on debt. These results are consistent with macroeconomic theory.

The transmission channel of the primary balance shows an ambiguous relation between risk and public debt, which calls for more detailed explanation: in the case of a primary surplus, a risk shock is followed by a nominal interest rate rise, as observed in equations (1) and (3). Consequently, as indicated in equation (6), the public debt increases.

Conversely, a primary deficit puts pressure on inflation. In this case, since inflation is already high because of the deficit, under the monetary policy rule (5), the interest rate is already high. Here, a risk shock will have a smaller effect on the interest rate and, thus, on public debt.

IV. Long-run equilibrium and the fiscal policy rule

The model's long-run dynamics can be configured on the basis of the results of the short-term comparative statics. This section manipulates the model to obtain a primary surplus rule for the fiscal authority to follow. Initially, the model presents three differential equations, namely:

$$\dot{R} = \sigma(r - \bar{i}) \quad , \sigma > 0$$

(2)

$$\frac{di}{dt} = \beta(\pi - \pi^*) + \alpha(b - b^*) + \mu(i^* - \bar{i}) \quad , \beta > 0; \alpha > 0; \mu < 0$$

(5)

$$\dot{b} = ib + g - t$$

(6)

By isolating *r* in equation (3) and substituting the result into (2), we observe the differential equation that gives default risk over time. We thus obtain:

$$\dot{R} = \sigma\left(i^* - \pi - \bar{i}\right)$$
(2.1)

The model equilibrium initially requires a system. However, a steady state dynamic is assumed for public debt such that the movement of debt over time can be taken as nil, i.e. the fiscal authority follows a rule under which it is passive and avoids incurring fiscal deficits. Thus, we obtain:

$$ib = g - t$$
(6*)

In other words, we assume a fiscal policy rule that gives a large enough primary surplus to cover public debt service, thereby keeping the debt sustainable over time, i.e. $\dot{b} = 0$. This manoeuvre allows the model, which previously had three differential functions, to be described as a system of two dynamic equations.

Accordingly, (2.1) and (5) may be recast as follows:

$$\dot{R} = \sigma\left[i^* - \bar{i} - \pi\left(R, i\right)\right]$$
(2*)

$$\frac{di}{dt} = \beta\left[\pi\left(R, i\right) - \pi^*\right] + \alpha\left[b\left(R, i\right) - b^*\right] + \mu\left[i^*\left(R, i\right) - \bar{i}\right]$$
(5*)

The system thus becomes a (2x2) one, in which equilibrium is drawn from equations (2.1) and (5). Thus, at equilibrium (steady state), we obtain:

$$\dot{R} = 0 \Rightarrow \pi\left(R, i\right) = i^* - \bar{i}$$

$$\frac{\partial i}{\partial t} = 0 \Rightarrow \pi\left(R, i\right) = \pi^* + \left(\frac{-\alpha\left[b\left(R, i\right) - b^*\right] - \mu\left[i\left(R, i\right)^* - \bar{i}\right]}{\beta}\right)$$

When the system is linearized around the equilibrium using a Taylor expansion, we obtain:

$$\frac{\partial R}{\partial t} = \sigma\left(-\frac{\partial \pi}{\partial R}\right)\left(R_* - R_0\right) + \sigma\left(i_* - i_0\right)$$
(12)

$$\frac{\partial i}{\partial t} = \left(\beta\frac{\partial \pi}{\partial R} + \alpha\frac{\partial b}{\partial R}\right)\left(R_* - R_0\right) + \mu\left(i_* - i_0\right)$$
(13)

Casting the results in matrix notation:

$$\begin{bmatrix} \dfrac{\partial R}{\partial t} \\ \dfrac{\partial i}{\partial t} \end{bmatrix} = \begin{bmatrix} \sigma\left(-\dfrac{\partial \pi}{\partial R}\right) & \sigma \\ \left(\beta\dfrac{\partial \pi}{\partial R} + \alpha\dfrac{\partial b}{\partial R}\right) & \mu \end{bmatrix} \begin{bmatrix} \left(R_* - R_0\right) \\ \left(i_* - i_0\right) \end{bmatrix}$$
(14)

The necessary and sufficient condition for the equilibrium of a two-dimensional dynamical system to be asymptotically stable (where the two eigenvalues of the solution have negative real parts) is that the trace and determinant of the Jacobian matrix be negative and positive, respectively.[8]

Thus, it is observed that:

$$Trace = \sigma\left(-\frac{\partial \pi}{\partial R}\right) + \mu = ?$$

$$Det = \sigma\left(-\frac{\partial \pi}{\partial R}\right)\mu - \sigma\left(\beta\frac{\partial \pi}{\partial R} + \alpha\frac{\partial b}{\partial R}\right) = ?$$

To satisfy the stability of equilibrium conditions, necessarily $|\mu| > \left|\sigma\left(-\frac{\partial \pi}{\partial R}\right)\right|$. Thus the trace will be negative.

In short, the first condition for stability is that, over time, the absolute value of the sensitivity of the nominal interest rate to the interest rate spread must be greater than the absolute value of the product of the impact of risk on inflation and the sensitivity of risk to the interest rate spread. There are thus two channels acting on macroeconomic stability: (i) the exchange-rate channel, which leads to stability, represented by μ, and (ii) the risk channel, which leads to instability, represented by $\sigma\left(-\frac{\partial \pi}{\partial R}\right)$.

This relationship is expected, given that, on the one hand, a reduction in the external interest rate pushes up the exchange rate, containing inflationary pressures occurring through that channel and thereby reducing the need to raise interest rates in the short term to meet the inflation target set by the monetary authority, in accordance with equation (5). On the other hand, a reduction in the external interest rate produces an effect on risk measured by σ, as given in equation (2). If risk is then high, this will require a significant rise in the domestic interest rate, as may be observed in (1), which will destabilize i. It should also be considered that if the negative impact of risk on inflation is very large,[9] then the rise in the interest rate may also be significant. As a consequence, for the trace $\sigma\left(-\frac{\partial \pi}{\partial R}\right) + \mu$ to be negative, then necessarily $|\mu| > \left|\sigma\left(-\frac{\partial \pi}{\partial R}\right)\right|$.

Another way of looking at this relationship is to admit the hypothesis that debt is directly correlated with GDP and σ, bearing in mind that this coefficient measures creditor mistrust in the government's ability to pay. Thus, public debt forms one of the channels through which an explosive effect on risk could be propagated, because the rise in interest pushes up the cost of debt servicing, increasing creditor mistrust of the government's ability to pay. A fiscal policy that takes public debt solvency into consideration could contribute to economic stability.

Given that fiscal policy was underlined as a possible instrument for achieving stability in the model, it is important to analyse possible channels of transmission with respect to the variables whose stability is studied. A surplus-targeting policy prevents the public debt from rising over time, thus stabilizing the default probability and, thus, the long-term interest rate (see diagram 3).

Diagram 3
Flow diagram 3: channel 1 of fiscal policy transmission

Source: Prepared by the authors.

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[8] See more detail in Gandolfo (1997).

[9]   It should be recalled that, according to the model, risk has a negative effect on inflation through the following mechanism: a rise in risk is accompanied by an increase in interest on government bonds, to ensure debt solvency, which pushes inflation down.

A second channel for the transmission of fiscal policy is that public spending containment cushions effects on aggregate demand. This eases inflation, which allows the monetary authority to lower the interest rate used to steer inflation towards the desired target, and to reduce government spending on debt servicing (see diagram 4). There are, then, two possible effects, as shown in the flow diagram below.

Diagram 4
Flow diagram 4: channel 2 of fiscal policy transmission

Source: Prepared by the authors.

Following analysis of the dynamic equilibrium stability, the determinant may be rewritten as follows:

$$Det = -\sigma \left[\frac{\partial \pi}{\partial R}(\mu + \beta) + \alpha \frac{\partial b}{\partial R} \right]$$

Thus, assuming a primary surplus and observing that $\frac{\partial b}{\partial R} > 0$, it may be deduced that $\beta > \mu$ is a necessary condition for the stability of the model's dynamic equilibrium. Since these coefficients indicate the sensitivity of the interest rate to the variation in inflation and in the interest rate spread, as expressed in (5), that condition is consistent with the expectations of the model. This is because controlling price levels is the premier aim in determining monetary policy instruments under an inflation-targeting regime.

In short, under a primary surplus fiscal rule, stable dynamic equilibrium requires that: (i) the base rate of interest be more sensitive to exchange-rate devaluation than to increases in risk; (ii) the fiscal set-up stabilize the confidence of external investors in public debt solvency, as reflected in σ, and (iii) the monetary authority concern itself more with deviations from target inflation than with the exchange rate.

For a qualitative analysis of the intertemporal trajectory and stability of the dynamic equilibrium, a phase diagram of the dynamic system expressed by equations (2) and (5) provides interesting study.

The system's convergence towards equilibrium can be studied by examination of the system discriminator, as given by: $D = [trace\,(A)]^2 - 4det(A)$, where D is the value of the discriminator and A is the matrix analysed.

When $D(A) > 0$, the equilibrium will be a stable node. Thus, in the event of a shock that leads to a deviation from equilibrium, the system will return monotonically to equilibrium. When $D(A) < 0$, the equilibrium will be a stable focus and the system will revert spirally to equilibrium after a shock.

In the model studied earlier, which represents the coordination of monetary and fiscal policies in the presence of a primary surplus rule, equilibrium is a stable focus and convergence thus occurs in a spiral wave form, where the derivative $\frac{\partial b}{\partial R} > 0$ is very low —and, thus, the primary surplus is small. When the surplus is large and, thus, the derivative is also very large, equilibrium could be a stable node with monotonic convergence.

Diagram 5 shows the scenario in which the system equilibrium is a stable node.

Diagram 5
Phase diagram: stable node
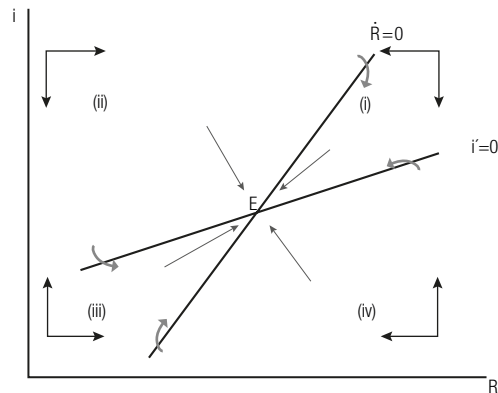
Source: Prepared by the authors.

Whether convergence occurs at a focus or node can have important implications in terms of economic policy implementation. This is because a combination of economic policies leading to convergence on equilibrium in a spiral or focal trajectory could weaken the credibility of monetary policy, insofar as the variables affected by economic policy decisions will necessarily register values above or below equilibrium at some points during the trajectory.

The phase diagram below represents a stable focus equilibrium (see diagram 6).

Diagram 6
Phase diagram: stable focus
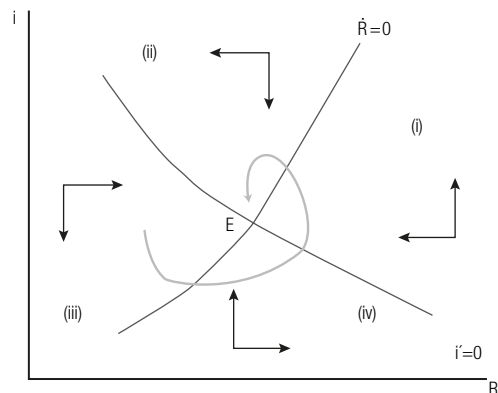
Source: Prepared by the authors.

V.   Extension of the model: the nominal exchange rate under the effects of default risk

Thus far, the exchange rate has been presented as a function of the interest rate spread. However, in an emerging economy, default risk is an important variable in determining the exchange rate and thus has an indirect effect on inflation. So, one way in which the central bank can lose control of monetary policy[10] is through the exchange rate which is, in turn, affected by risk.

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[10] Loss of control of monetary policy is understood to mean a situation in which the monetary authority cannot meet its preset inflation target.

In light of these considerations, there follows a study of the comparative statics of the economic model in a situation where the exchange rate is influenced by default risk. Equation (9), which denotes the exchange rate (strictly speaking, the devaluation of the exchange rate), is represented as follows:

$$E = \rho\left(i^* - \bar{i}\right) + \gamma R \ , \ \ \rho < 0; \gamma > 0$$

(9.1)

where ρ measures the exchange rate's sensitivity to the difference between the domestic and external interest rates and γ measures the exchange rate's sensitivity to risk. This parameter, in turn, can also be considered a measure of risk aversion on the part of international investors, like σ, since at a particular risk level, the larger γ, the less willing investors will be to add the respective country's bonds to their portfolio. This shrinks the supply of foreign exchange and increases exchange-rate depreciation. Following Blanchard (2004), it may also be expected that γ will rise if the debt-to-GDP ratio increases.

Blanchard (2004) empirically estimated a similar function to the one discussed and found that risk had the expected effect —and of a large magnitude— on the exchange rate. Favero and Giavazzi (2003) found similar results.

Let us consider the Phillips curve presented in equation (7.1):

$$\pi = \tau\left(y - \bar{y}\right) + \theta\left(E\right)$$

(7.1)

By substituting the IS curve presented in (10.1) into the Phillips curve (7.1) and substituting (1) and (4) into (3) —to obtain $i^* = \left(r^e + R\right) + \pi$— and then inserting the results into the Phillips curve (7.1), along with the exchange-rate function (9.1), we obtain the following for inflation after reorganization:

$$\pi = \left[\frac{\left(\frac{\tau I_i}{1-c_y}\right)}{v} + \frac{\left(\theta + \frac{\tau X_e}{1-c_y}\right)}{v}\left(\rho + \gamma\right)\right]R + \left[\frac{\left(\frac{\tau I_i}{1-c_y}\right)}{v} + \frac{\left(\theta + \frac{\tau X_e}{1-c_y}\right)}{v}\rho\right]r^e$$
$$- \frac{\left(\theta + \frac{\tau X_e}{1-c_y}\right)}{v}\bar{i}\rho + \frac{\left(\frac{\tau}{1-c_y}\right)}{v}g - \frac{\tau}{v}\bar{Y}$$

(7.4)

where $v = \left[1 - \left(\frac{\tau I_i}{1-c_y}\right) + \left(\theta + \frac{\tau X_e}{1-c_y}\right)\rho\right] > 1$

From which the following partial derivative may be extracted:

$$\frac{\partial\pi}{\partial R} = \frac{\left(\frac{\tau I_i}{1-c_y}\right)}{v} + \frac{\left(\theta + \frac{\tau X_e}{1-c_y}\right)}{v}\left(\rho + \gamma\right)$$

(7.3.2)

There is some ambiguity in the sign of this partial derivative, because it should be negative in normal conditions —as shown in (7.3.1)— but positive in extreme cases where γ —which represents risk elasticity to the exchange rate— is too high. In other words, the derivative will have the opposite sign to that expressed in (7.3.1), if $\gamma > \rho$. This is because at high risk aversions, a rise in the interest rate

may not be enough to contain inflation, since capital flight can push prices up through the mechanism of exchange-rate pass-through.

Similarly, the exchange-rate effect given in (9.1) is plugged into the government's intertemporal constraint, represented by (6.3). Thus, considering that the government's budgetary constraint is given by:

$$g - t = \left(r^e + R + \pi\right)b \tag{6.3}$$

We insert (7.4) into (6.3) and solve for the level of public debt, such that

$$b = \frac{g - t}{\left\{\left[1 + \dfrac{\left(\dfrac{\tau I_i}{1 - c_y}\right)}{v} + \dfrac{\left(\theta + \dfrac{\tau X_e}{1 - c_y}\right)}{v}(\rho + \gamma)\right]R + \left[1 + \dfrac{\left(\dfrac{\tau I_i}{1 - c_y}\right)}{v} + \dfrac{\left(\theta + \dfrac{\tau X_e}{1 - c_y}\right)}{v}\rho\right]r^e + \dfrac{\left(\dfrac{\tau}{1 - c_y}\right)}{v}g - \dfrac{\left(\theta + \dfrac{\tau X_e}{1 - c_y}\right)}{v}i\rho - \dfrac{\tau}{v}\bar{Y}\right\}} \tag{6.4}$$

To analyse the effect of a marginal rise in risk on the behaviour of public debt, we extract the following partial derivative:

$$\frac{\partial b}{\partial R} = \frac{(t - g)\left[1 + \dfrac{\left(\dfrac{\tau I_i}{1 - c_y}\right)}{v} + \dfrac{\left(\theta + \dfrac{\tau X_e}{1 - c_y}\right)}{v}(\rho + \gamma)\right]}{i^2} > 0 \tag{6.3.2}$$

> 0 for $t>g$ or, < 0 for $t<g$.

This derivative, which indicates the risk elasticity of public debt, will be positive in the case of a primary surplus and negative in the case of a deficit, as described in section III.2.

However, it should be noted that once risk is included as an explanatory variable for the exchange rate, unforeseen effects may be anticipated in the previous configuration of the model. That is, if $\gamma>\rho$, the derivative may again have a different sign than expected. Following the example given in IV.1, deductions can be made about possible changes in the long-term stability of the equilibrium for the model with a primary surplus fiscal rule, this time including the hypothesis of risk as an explanatory variable for the exchange rate.

Although the model matrix remains identical to that presented in (14) under section IV.1, now —with a nominal exchange rate influenced by risk and, thus, possible alterations in the direction of the partial derivatives— results other than those expected may occur.

The stability of the model will depend, as noted earlier, on the signs of the trace and determinant of the matrix. These also remain the same as those presented before, with the exception of possible changes in sign generated by the impact of risk on the exchange rate. Thus:

$$Trace = \sigma\left(-\frac{\partial \pi}{\partial R}\right) + \mu = ?$$

$$Det = -\sigma\left[\frac{\partial \pi}{\partial R}(\mu + \beta) + \alpha\frac{\partial b}{\partial R}\right]$$

The difference between this configuration and the previous one is that now stability is influenced by the effect of risk on the exchange rate. Even with a primary surplus, which supports an asymptotically stable equilibrium in the first model, too high a risk aversion (as measured by γ) can lead to instability in the event of an exogenous shock generated by the rise in risk itself. In other words, even assuming all the necessary conditions for stability —as in section IV.1—, the Jacobian determinant may be negative if $\gamma > \rho$, because then the partial derivatives that measure the impact of risk on inflation and public debt may present unexpected signs. As seen earlier, at very high levels of risk aversion, a risk shock can lead to an exchange-rate devaluation that pushes up inflation and induces instability in the model.

VI. Analysis of findings and concluding remarks

This article has analysed the coordination of monetary and fiscal policies in an emerging economy with an inflation-targeting regime, in a context in which default risk shocks can lead to macroeconomic disequilibria. It has sought to understand how the economy adapts to exogenous shocks to maintain an asymptotically stable equilibrium.

On the basis of a model proposing a monetary policy rule that takes into account not only the deviation of inflation from its target $(\pi - \pi^*)$, but also the deviation of public debt from its desired level $(b - b^*)$ and the spread between domestic and external interest rates $(i^* - \bar{i})$, we identified the model's equilibrium stability conditions, which are summarized below. The form used was a model of simultaneous first-order differential equations, analysing their intertemporal equilibrium and their stability. It was concluded that:

(a)   For a model in which the exchange rate is defined by interest rate parity and a primary surplus regime, stability requires that: (i) monetary policy afford more importance to the deviation of inflation from its target than to the interest rate spread; (ii) shifts in the interest rate spread have a greater influence on monetary policy than the need to adapt short-term interest rates to risk shocks. It may be deduced from these conditions that controlling price levels should be the chief concern of monetary policy in an inflation-targeting regime and should thus be afforded more importance that the need to adapt to external interest rate shocks. At the same time, fiscal oversight is needed to make the economy less vulnerable to default risk shocks.

(b)   Under a scheme of policy coordination in which the nominal exchange rate is defined by the interest rate spread and by the risk factor, a very high level of risk aversion, as measured by γ, could have a large enough impact on the nominal exchange rate to destabilize the model. This is because a rise in interest could be interpreted as a greater probability of default and could consequently lead to exchange-rate devaluation and, potentially, loss of monetary policy control over inflation. For this reason, in economic policy terms, a policy of fiscal austerity is recommended. The other conclusions mentioned remain valid under this configuration. These results are similar to those obtained by Blanchard (2004) and Favero and Giavazzi (2003).

Generally speaking, the results suggest that inflation control should be the monetary authority's main objective in an inflation-targeting regime. At the same time, fiscal policy should be passive, generating surpluses to stabilize the public debt and ensure its solvency over time, which will, in turn, stabilize the default risk and avoid the risk of fiscal dominance.

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Mergers and acquisitions carried out by Spanish firms in Latin America: a network analysis study[1]

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Abstract

This article analyses the changing role played by Spanish firms in foreign investment through mergers and acquisitions in Latin America in 1999-2012. Spanish enterprises that acquired local assets in Latin America and became leading players in the late 1990s have seen the situation change drastically, as new competitors have emerged to undermine Spain's importance as an investor in the region. This study uses the social networks methodology, which enables a more complex analysis than traditional approaches, by studying the position of agents as members of a network being analysed. The use of centrality, density and centralization indicators reveals the structure of the network and how it changes through time, and thus, it shows the relative position of each investor country, Spain in particular.

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I.   Introduction

Foreign direct investment (FDI) has become one of the core elements of the process of globalization and structural changes in the world economy. This has had a variety of repercussions, such as a redefinition of the new geographical dimensions of political economy, the formation of new power relations between economic agents in the national and international spheres, the increased importance of flows of financial and human capital between countries, or technology transfer. Despite the importance of developed countries as both origins and destinations of investments, developing economies have grown in importance over the last few years. Whereas in 1990, the advanced economies generated 95.3% of FDI outflows and received 83.1% of the inflows, over the last 10 years, the developing countries have become the main recipients of foreign investment (60.8% of the total in 2013), and their firms are becoming major international investors, accounting for 39% of total FDI (UNCTAD, undated). This is largely explained by the success of the emerging economies and the decline of the triad formed by the United States, the European Union and Japan. Global transformations are not only occurring in the strictly economic domain, but also in other spaces of power such as the technological, political-diplomatic and military spheres, among others. This has repercussions in both the public (States) and the private spheres (transnationals), raising new problems for the governance of national economies, as summarized by Rodrik (2011) in what he refers to as the political trilemma of the world economy.

This is the international backdrop facing the Spanish economy and its firms, which launched their internationalization process in the mid-1990s and was concentrated geographically in Latin America and sectorally in telecommunications, energy and finance. This is explained by a coincidence in time between two factors: (i) the need for Spanish firms to internationalize their operations as a result of the greater competition fuelled by the deepening of the European internal market (1992) and the creation of the Economic and Monetary Union (1999), and (ii) the financing needs of the Latin American economies to implement the structural reforms imposed following the external debt crisis of the 1980s and 1990s (Sánchez Díez, 2002). The changes that have occurred in the Spanish economy, particularly the privatization of public enterprises and external liberalization, strengthened the ownership advantages of some large national firms; while the opening up of capital accounts, the deregulation of regulated sectors, and privatization processes in many Latin American countries improved that region's location advantages, as described by Dunning (1988 and 1994).

Today, the context in which Spanish firms are investing in Latin America is very different from what they encountered two decades ago, owing to the changes that have occurred both in Spain and in that region. Spain is going through a profound crisis that is having undeniable repercussions on the behaviour of national firms. Although they have not decided to disinvest abroad on a massive scale, they have been forced to reorganize their international assets. Moreover, Latin America has achieved high rates of economic growth, sustained by the dynamism of domestic demand and that of emerging countries, in a framework of macroeconomic stability and consolidated public-sector accounts. Accordingly, many countries have designed more sophisticated policies to attract foreign investment, prioritizing "quality" over "quantity," while some Latin American firms have themselves started to internationalize.

In view of this new scenario, this article analyses the investor status of Spanish firms in Latin America and how this has altered as a result of the changes described above. The countries whose investments are analysed are: Argentina, the Bolivarian Republic of Venezuela, Brazil, Chile, Colombia, Costa Rica, Cuba, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Nicaragua, Panama, Paraguay, Peru, the Plurinational State of Bolivia and Uruguay, within the region; and China, France, Germany, Italy, Japan, the United Kingdom, the United States and the countries of the BENELUX economic union (Belgium, the Netherlands and Luxembourg), outside the region.

The aims of this article are as follows:

- Aim 1: to study the dynamics of business mergers and acquisitions in Latin America in 1999-2012, particularly in terms of the degree of interaction between countries and their geographical concentration.

- Aim 2: to evaluate the role played by Spanish firms in mergers and acquisitions in Latin America, compared to their potential competitors, observing their trend over the period studied. This will make it possible to examine whether or not Spain is losing the leadership position it attained in the second half of the 1990s. In terms of network analysis, the study will detect whether a shift is occurring from central positions to more peripheral ones.

- Aim 3: to analyse the dynamics of mergers and acquisitions in the telecommunications, energy and finance sectors. The aim here is to elucidate whether there are significant differences arising from the different economic activities and identify the position held by Spanish firms in each sector.

To address these aims, the social network analysis methodology was chosen for its great capacity to explain economic independencies, as evidenced in several previous research studies. Network analysis makes it possible to systemically study the complex dynamics of interaction, influences and interdependencies between countries, going beyond the financial data on business mergers and acquisitions in Latin America.

Following this Introduction, section II reviews theoretical work on foreign direct investment and the application of social network analysis to the study of international economic relations. Section III describes the key methodological aspects of the construction, interpretation and analysis of merger and acquisition networks. Section IV analyses the network of mergers and acquisitions among firms in Latin America and the position occupied by Spain compared to its competitors, presenting a dynamic study of 1999-2012, and a static study of the telecommunications, energy and finance sectors. Lastly, section V sets out the conclusions of this article.

II. Theoretical aspects of mergers and acquisitions and social network analysis

Dunning (1977, 1979, 1980 and 1988) summarized the contributions made by different authors to the study of FDI in the so-called "eclectic" paradigm, indicating that FDI would take place in the presence of advantages in terms of ownership, the location of the destination country[2] and internalization (which is known as the OLI model or paradigm). From the industrial organization school, he drew on contributions by Hymer (1976 and 1979), Kindleberger (1969), Caves (1971, 1980 and 1982) and Caves and Hirschey (1981), which gave rise to "ownership advantages," in other words specific capacities that multinational enterprises have in terms of exclusive ownership compared with local firms. Nonetheless, while necessary, ownership advantages are not sufficient for a company to invest abroad. This requires the presence of a location advantage in the destination country, an element that is analysed by location theories. The contributions of internalization theory were included as well, specifically those by Williamson (1975), Casson (1979 and 1985), Rugman (1976, 1980 and 1981) and Tecee (1986), who started from the hypothesis that enterprise internalization is a mechanism for reducing the transaction costs arising from imperfectly functioning markets.

The criticisms that have been levelled against the OLI paradigm include the fact that it cannot clearly explain investment originating in developing countries, since most of those countries' firms do

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not have exclusive ownership advantages. According to Moon and Roehl (2001), the transnational enterprises of emerging countries pursue their own strengthening with their internationalization, by accumulating resources and assets to which they previously did not have access. Even Dunning himself (2009) has nuanced his theoretical reflections, by arguing that the pursuit of knowledge and learning experiences could be motivating internationalization among firms from those countries, in the absence of the classic ownership advantages —an idea previously put forward by Luo and Tung (2007). In short, the international expansion of developing-country firms can be interpreted in the light of the analytical framework known as linkage, leverage and learning (the LLL model) (ECLAC, 2014; Mathews, 2006), whereby the firms make a variety of alliances to gain access to resources that were previously unavailable to them. Thus, firms that start later in the internationalization process, can take advantage of other firms' experience and knowledge.

Nonetheless, it is clear that the reception of foreign firms represents an opportunity for domestic firms to appropriate the externalities that foreign investments generate in the destination economies. This can improve their location attraction for future investments, as argued in the OLI paradigm; but it can also encourage domestic enterprises to embark on internationalization strategies themselves, as suggested by the LLL model.

For these reasons, and despite this being a controversial topic,[3] two assumptions are made:

(i)   The presence of transnational enterprises in a given economy can result in changes to the production fabric, through spillover effects (Aitken and Harrison, 1999): a merger or acquisition can result in certain ownership advantages being transferred to the acquired local asset. These transfers are understood as flows of information, knowledge, learning on modes of organization, human capital, financing and technology. The economy acquiring assets abroad can also receive spillover effects in the form of knowledge flows, new business practices or human resources.

(ii)   Transnational enterprises can encourage domestic firms to modernize their production and training processes and upgrade the quality of their production (Blomström and Kokko, 2003). Owing to spillover effects, the two countries participating in a cross-border operation can obtain specific benefits from these relations with an impact on their economic structures.[4]

Given that countries, through their firms, tend to engage in merger and acquisition operations with several countries, the aforementioned spillover effects ultimately occur multilaterally, as shown in diagram 1.

Diagram 1
Example of multilateral spillovers in cross-border merger and acquisition operations
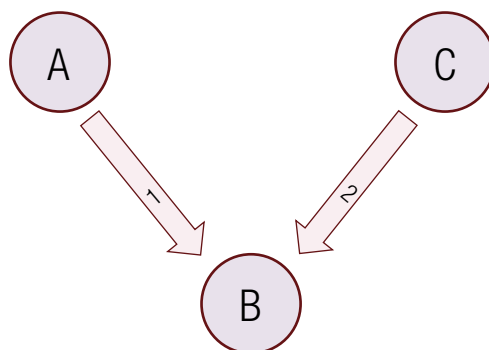
Source: Prepared by the authors.

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[3]   For a summary of positions on this topic, see box I.1 of ECLAC (2011, p. 29).

[4]   The dissemination of the benefits and their structural impact will not be automatic or necessarily direct; and the benefits will not necessarily occur in all merger or acquisition operations. Public policies are decisive in determining the final outcomes.

Link 1 represents acquisitions made by firms in country A of businesses in country B, which involves a foreign direct investment flow from A to B. Link 2 reflects the same process between countries B and C. The operations represented in link 1 can generate transformations in countries A and B; but these could also occur in country C through Link 2. In other words, spillovers of information, knowledge, organizational learning, human capital, financing and technology are not exclusively confined to direct links between two countries, but circulate between them through multiple interactions, under a network logic. It is therefore essential to simultaneously consider the set of relations that exist between countries to gain an adequate understanding of the flows in question.

As noted above, cross-border mergers and acquisitions are analysed using the social networks methodology, so the characteristics of the individual social units will be viewed as arising out of structural or relational processes (Wasserman and Faust, 1994, p. 7). Broadly speaking, studies on this methodology include theoretical analyses of the formation and evolution of networks (Watts and Strogatz, 1998; Jackson and Wolinsky, 1996), empirical studies on the structures and patterns of networks in the real world (Bearman, Moody and Stovel, 2004; Fagiolo, Reyes and Schiavo, 2009) and methodological analyses that provide new research tools (Wasserman and Faust, 1994; Jackson, 2008).

In economics, this methodology has been used to analyse the structure and functioning of competitive markets (Mitchell and Skrzypacz, 2006; Amir and Lazzati, 2011), employment and wage inequality (Calvó-Armengol and Jackson, 2004), the dissemination of information and innovations (Schilling and Phelps, 2007; Fleming, King and Juda, 2007; Galaso, 2011) and the patterns governing international trade (Kali and Reyes, 2006; De Benedictis and others, 2013).

There is also a growing academic literature that uses network analyses to study international financial flows in general, although its application to the analysis of foreign investment is still relatively incipient. Some studies have focused on analysing international financial crises (Chinazzi and others, 2013; Elliott, Golub and Jackson, 2014), and the shareholding structures of transnational enterprises (Vitali, Glattfelder and Battiston, 2011; Vitali and Battiston, 2013). In contrast, Haberly and Wojcik (2013) focus on the FDI network originating in tax havens, and show a heavy concentration of flows and high level of dependency on political and social similarities between countries, while Visintin (2011) analyses the international networks of trade and FDI, which display a hub-and-spoke structure.

III. Methodology: construction and analysis of merger and acquisition networks

The following paragraphs detail the methodological aspects of constructing networks applied to the case of cross-border mergers and acquisitions, and the analysis of the results obtained.

1. Construction of the networks

In this case, applying this methodology requires a double-entry matrix of the economic flows that took place between pairs of countries in 1999-2012. As there is no information at the global level on FDI by origin and destination, or by sector, for all countries, it was decided to construct the matrix from the database on cross-border mergers and acquisitions prepared by Thomson Reuters. The following have been calculated: 1 matrix and 1 network covering the entire period and all sectors; 7 biennial matrices and 7 biennial networks (except for one case in which they are triennial and which corresponds to just one year, the end of the period); and 3 sectoral matrices and 3 sectoral networks. In other words, a total of 11 matrices and 11 networks have been calculated, along with their associated indicators.

The entire network requires two basic elements: nodes and links. In this analysis, the nodes correspond to the 28 national economies covered by the study, as mentioned above; and the links between the nodes represent the total value (in millions of dollars) of the purchases made by the firms of one country of those in another. From the database used, all operations were selected in which at least one of the two firms involved —either the buyer or the seller— is located in a Latin American country. In other words, relations between two non-Latin American economies were not included.[5]

Table 1 shows a simplified example of the database, with three operations between firms belonging to two countries. By aggregating these operations, the values of the links connecting the countries are calculated as shown in diagram 2.

Table 1
Sample of the database used to construct the networks of business cross-border mergers and acquisitions

Buyer		Seller		Size of the operation
Firm	Country	Firm	Country	
1	A	2	B	100
3	A	4	B	50
5	B	6	A	200

Source: Prepared by the authors.

Diagram 2
Example of two links between countries in business merger and acquisition operations
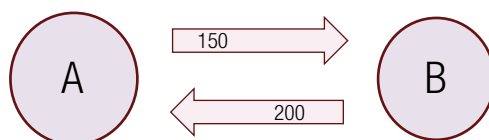
Source: Prepared by the authors.

Once the links are identified, their direction and weights are analysed. The direction depends on where the buying and selling firms are located. Diagram 2 shows a link that runs from A to B and another that goes from B to A. Having directed links makes it possible to distinguish between inflows and outflows, and this broadens the options for calculating and interpreting the network indicators. Accordingly, this is a directed network or graph, in which the links can be analysed in both directions. The weight corresponds to the total value of the purchases made by the firms of one country of those in another. Thus, in diagram 2, the weight of the link running from A to B is 150, representing the sum of mergers between firms 1 and 2, on the one hand, and 3 and 4, on the other. Similarly, the link that runs from B to A has a value of 200, equivalent to the merger between firms 5 and 6. This makes it possible to identify the most important links in the network and those that are less important. Simultaneously tracing all of the links between the 28 countries defines the full network of business mergers and acquisitions in Latin America.

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[5]   For example, this analysis includes a purchase of Colombian asset by a French firm, and vice versa, as well as the acquisition of a Chilean company by a Brazilian firm; but it excludes operations between French and Spanish enterprises.

2.   Network analysis

To study the dynamics of enterprise mergers and acquisitions in Latin America (aim 1) the full network is analysed, which includes data for all years and the seven time-framed networks. These networks are studied from a global perspective, that is, using indicators that describe the structure of the network as a whole, without distinguishing the position occupied by each country. Specifically, it uses indicators of density and total degree centrality, which measure structural properties relating to the patterns of collective interaction between all of the countries included. These indicators are interpreted as follows:

•   Density measures the proportion of links existing in a network relative to the maximum possible number,[6] and thus quantifies the degree of connection between actors. In this study, it is used to measure the overall level of interaction between countries.

•   The total degree centrality indicator measures concentration in the distribution of links between network nodes, in other words the concentration of merger and acquisition operations between countries. This makes it possible to quantify whether these operations are concentrated among just a few countries that channel resources, accumulate spillover effects and exert their influence on the rest; or, on the contrary, whether the merger and acquisition links are established in a relatively distributed way, with the participation of multiple centres of action.

The role played by Spain is then analysed, along with its evolution during the period studied, both in the full network of data from all years and in the seven time-framed networks (aim 2). The study is conducted using indicators that describe the relative position occupied by each country in the networks. In particular, four measures of centrality are calculated: in-degree, out-degree, total degree, and eigenvector. These are defined as follows:

•   Degree centrality corresponds to the number of links that a node has, weighted by the value of each link. In other words, it is the number of connections a country maintains with the other countries in the network, weighted by the value, in millions of dollars, of the operations that these connections represent. In directed graphs (such as those used in this analysis) three different measures of degree centrality can be defined: in-degree, out-degree and total. Their calculation is analogous, but it is done by considering the links flowing into the node, those flowing out of it, and the sum of the two, respectively. These represent the volume of investment received by the country through sales of its firms to foreign companies (in-degree centrality); the outward investment by the country's firms in payment for assets purchased abroad (out-degree centrality); and the combination of the two (total degree centrality).

•   Eigenvector centrality (Bonacich, 1972) measures the influence capacity of a node, in this case a country. Nodes (countries) with high values of this indicator are linked to others which are also well connected, so they are in a position to exert influence, disseminate information or propagate spillover effects.

Lastly, the aim of studying the telecommunications, energy and finance sectors (aim 3) requires the use of sectoral networks to calculate all of the foregoing indicators. This makes it possible to analyse both the patterns of collective interaction between countries and the relative position of Spain in the three selected sectors.

Table 2 provides greater detail of the relation between the aims of this study, the associated economic concepts, the network indicators to be used for the analysis, and how they are calculated.

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6   The maximum number of links is attained when at least one firm in each of the 28 countries considered purchases a firm in each of the other 27 countries.

Table 2
Main network indicators proposed for the analysis

Aims of the article	Associated economic concept	Network indicator[a]	Calculation[b]
1 and 3	Interaction between the countries in business merger and acquisition processes	Density	$$D = \frac{V}{max(V)}$$ where V is the number of links existing in the network and $max(V)$ is the maximum number of links that could exist in the network if all nodes were linked to all others.
	Geographical concentration of business mergers and acquisitions	Total degree centrality	$$CGT = \frac{\sum_{u=1}^{u=|N|} CG_{tot}(u^*) - CG_{tot}(u)}{max\left(\sum_{u=1}^{u=|N|} CG_{tot}(u^*) - CG_{tot}(u)\right)}$$ where N is the number of nodes in the network u represents a given network node, $CG_{tot}(u)$ is the total degree centrality of node u and $CG_{tot}(u^*)$ is the maximum total degree centrality registered by a network node.
2 and 3	Relative importance of a country in the region's business mergers and acquisitions	Total degree centrality	$$CG_{tot}(u) = \sum_{v=1, v \neq u}^{|N|} w_{v,u} + \sum_{v=1, v \neq u}^{|N|} w_{u,v}$$ where N is the number of nodes in the network, v represents a given node of the network, $W_{v,u}$ is the value of the link running from node v to node u and $W_{u,v}$ is the value of the link running from node u to node v.
		Eigenvector centrality	$$\lambda CVP = W \cdot CVP$$ where W is the adjacency matrix, i.e. the square matrix that represents all network links, and λ is the maximum eigenvalue of the adjacency matrix.
	Relative importance of a country in the region's mergers and acquisitions (by purchasing firms)	Out-degree centrality	$$CG_{sal}(u) = \sum_{v=1, v \neq u}^{|N|} w_{u,v}$$ where N is the number of nodes in the network, v represents a given node of the network, and $W_{u,v}$ is the value of the link running from node u to node v.
	Relative importance of a country in the region's mergers and acquisitions (by firms purchased)	In-degree centrality	$$CG_{ent}(u) = \sum_{v=1, v \neq u}^{|N|} w_{v,u}$$ where N is the number of nodes in the network, v represents a given node of the network, and $W_{v,u}$ is the value of the link running from node v to node u.

Source: Prepared by the authors.

[a]   The density and centralization indicators reveal general characteristics of the network, whereas the degree centrality indicators (total, in and out) and the eigenvector centrality indicator make it possible to observe the position of a node (country) in the network.

[b]   The centrality indicators are normalized by dividing them by the maximum possible value.

IV. Results of the network analysis on mergers and acquisitions in Latin America

The following paragraphs report the results of the network analysis of mergers and acquisitions among firms in Latin America between 1999 and 2012. Firstly, the general dynamics of mergers and acquisitions in the region are studied; later, the role played by Spain is analysed, and lastly, the analysis focuses on the telecommunications, energy and finance sectors.

1. The dynamic of mergers and acquisitions in Latin America

The structural properties of the global network are firstly analysed for the years studied (see figure 1 and table 3), focusing on the degree of interaction between countries (density) and the geographical concentration of this type of operation in the region (centralization). The results reveal two key features:

- The degree of interaction between countries, measured by the network density indicator, is highest in two periods, 2007-2009 and 2010-2011. This reflects a major change in the reorganization of productive assets in Latin America. In the late 1990s and early 2000s, business acquisitions in the region were undertaken by companies from just a few countries. Spanish firms, in particular, took advantage of the privatization processes that were unfolding in the vast majority of Latin American countries. Nonetheless, the reality of the region changed rapidly. The strengthening of Latin American firms and the boom in their operations abroad have increased the degree of interaction between the countries of the region, particularly considering that the vast majority of trans-Latin enterprises have their foreign assets located elsewhere in Latin America itself. Chinese firms are also displaying increasing interest, with a very aggressive dynamic in terms of natural resource extraction; while United States companies have always been present in the region, albeit in cycles.

- In terms of the geographical concentration of mergers and acquisitions, the total centralization indicator is at a maximum in 2007-2009, which shows that the large number of merger operations, revealed in the increase in density, progressively tended to concentrate in fewer countries. Moreover, as from 2010, this indicator remains above the levels recorded in the first half of the 2000 decade, which shows that merger and acquisition opportunities since the international financial crisis are being exploited by fewer countries. In other words, this reflects a trend towards a network structure in which most countries are relatively marginalized from cross-border business merger and acquisition operations, while just a few account for most of the investment inflows and outflows.

Figure 1
Latin America: network of cross-border business mergers and acquisitions, 1999-2012
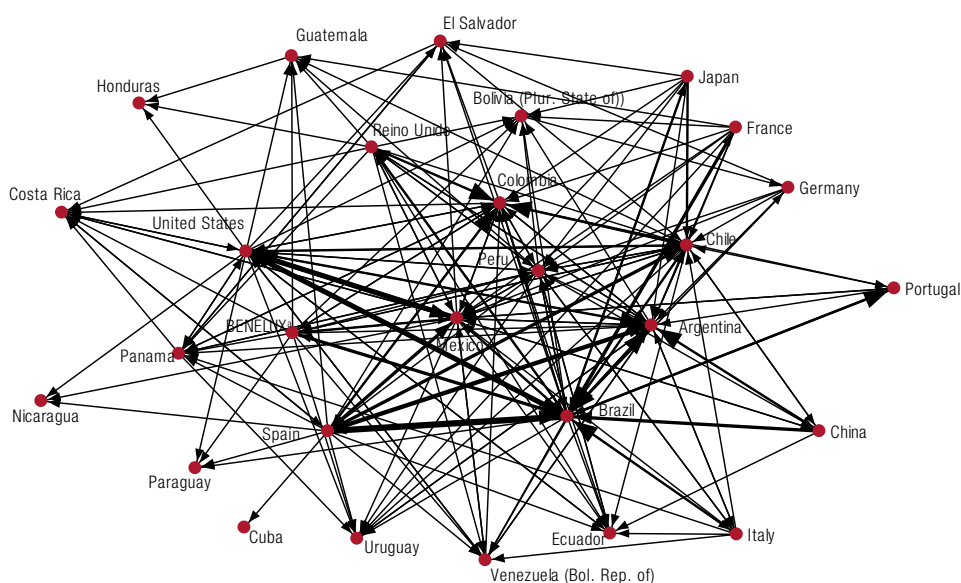
Source: Prepared by the authors on the basis of data from Thomson Reuters.

a Economic union between Belgium, the Netherlands and Luxembourg.

	1999-2000	2001-2002	2003-2004	2005-2006	2007-2009	2010-2011	2012
Density	0.118	0.101	0.091	0.106	0.159	0.143	0.081
Total degree centrality	0.044	0.03	0.038	0.050	0.106	0.073	0.051

Source: Prepared by the authors.

2.   The role of Spain in mergers and acquisitions in Latin America

The analysis now focuses on the role played by Spain in the network of business mergers and acquisitions in Latin America, by studying its position relative to the other countries (nodes) of the network. Special attention is paid to the comparison with countries that have been very active in recent years, such as Brazil, China and the United States.

The total degree centrality indicators by country (see figure 2) show that there is a small group of countries with high levels of centrality, and a large number of countries that occupy more peripheral positions —in other words, countries with weaker merger and acquisition flows (links) with the rest. The situation of the most central countries shows that their firms have established contacts, through mergers and acquisitions, with firms in a wide range of countries, whether as buyers (out-degree centrality) or as bought (in-degree centrality).

Figure 2
Selected countries: degree centrality indicators in cross-border business mergers
and acquisitions, 1999-2012
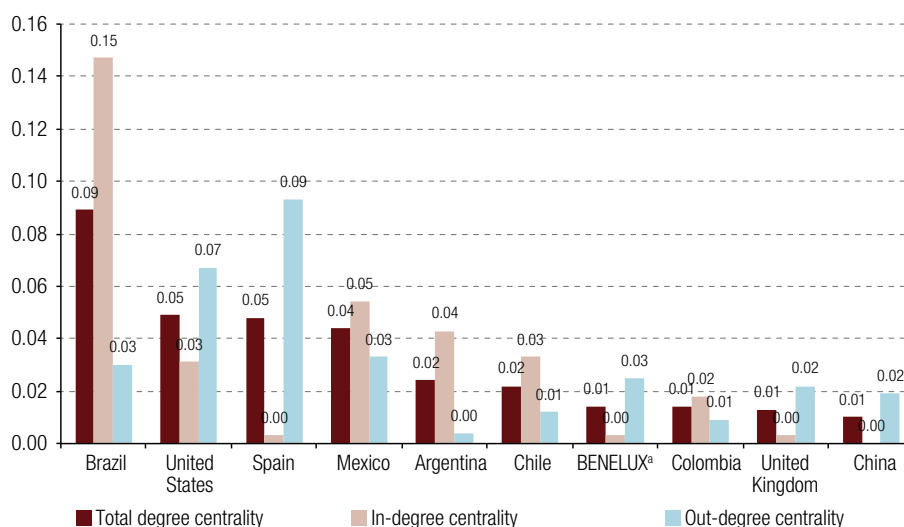
Source: Prepared by the authors.
a Economic union of Belgium, the Netherlands and Luxembourg.

In the entire period studied, the leading players are Brazil, Spain, the United States and, to a lesser extent, Argentina, Chile and Mexico. These are the economies which, from a global standpoint, are at the centre of the network (according to the total degree centrality indicator) maintaining most interactions with all agents participating in the network. There are significant differences between

them, however, particularly in terms of their differentiated position as countries of origin of the firms buying or selling assets. Given the assumptions, these economies are the most important in terms of their potential to channel spillover effects (Aitken and Harrison, 1999) —in other words their potential to facilitate flows of information, knowledge, organizational learning, human capital, financing and technology. They also display greater possibilities for modernizing their production processes, training and location in production (Blomström and Kokko, 2003), and for generating changes in their economic structures.

Spain and the United States, followed by Brazil, Mexico and the BENELUX countries (Belgium, the Netherlands and Luxembourg) are more dynamic in terms of the internationalization of their firms, as shown by their higher out-degree centrality indices. Spain's central role stems from its seizing the purchase opportunities generated by the 1990s privatization processes, and in the subsequent reorganization of assets and expansion of its companies' investments in the region, to consolidate their leadership and exploit a market that was already known to them. Although the entrepreneurial drive of Spanish companies has never faded, it was less intensive in the periods 2001-2002 and 2005-2006, when these firms retargeted their foreign operations towards the European market, while taking advantage of the experience and size they had acquired in Latin America. Spanish firms have to some extent reproduced their parent companies in their Latin American branches, which have been created through local takeovers (ECLAC, 2003 and 2012).

The United States, which was a major investor throughout the twentieth century, has continued to make large-scale purchases of productive assets in Latin America, as shown by its levels of out-degree centrality, which rose in the two years leading up to the crisis. This indicator has risen substantially also in the case of Brazil, reflecting the boom in the international expansion of its firms since the middle of the 2000 decade (ECLAC, 2005 and 2014). Brazilian trans-Latin companies are favoured by the substantial support policy applied by the public sector, in particular that carried out by the Brazilian Development Bank (BNDES). For its part, Mexico has held positions of greater out-degree centrality in the network, particularly in the period prior to the 2009 recession.

In terms of a country's attractiveness as an investment destination, Brazil and Mexico, and to a lesser extent Argentina and Chile, are the leaders. These economies have implemented far-reaching structural reforms that have made it easier for foreign firms to enter their economies and be profitable. In addition, both Brazil and Mexico have large domestic markets, a fact that appears to be an important location advantage for foreign investors. Nonetheless, the differences over the period are significant. Argentina's position deteriorated sharply, from high in-degree centrality in the late 1990s to very low levels in the latter years of the period analysed. The uncertainty triggered by the end of convertibility in 2001, which resulted in numerous disputes being brought before the International Centre for Settlement of Investment Disputes (ICSID) (Stanley, 2004; Zabalo Arena, 2012), caused a fall in centrality levels from which the country has not yet recovered during the twenty-first century (see figure 2).

The relatively low levels of in-degree centrality displayed by the United States and Spain are not surprising, because the way the network has been constructed means that only the acquisitions made by Latin American firms of Spanish or United States companies are included. This shows that trans-Latin firms are still relatively inactive in terms of expanding outside the region.

Another approach is to analyse a country's degree of connectedness with the most influential actors in the network, through eigenvector centrality. Brazil and Spain, followed by the United States, Mexico and Argentina, are the countries with the highest eigenvector centrality, since the transfer of information and knowledge, changes in modes of business organization, human resources, availability of information and, in general, the concentration of power revolve around those countries (see figure 3).

An analysis through time highlights the progressive loss of influence of countries such as Mexico or the United Kingdom, as measured by eigenvector centrality. Another key feature is the sudden emergence of China which, having been absent throughout the period, in 2010-2011 achieved a centrality level ahead of Spain, to rank second behind Brazil.

Figure 3
Selected countries: eigenvector centrality in cross-border business mergers and acquisitions, 1999-2012
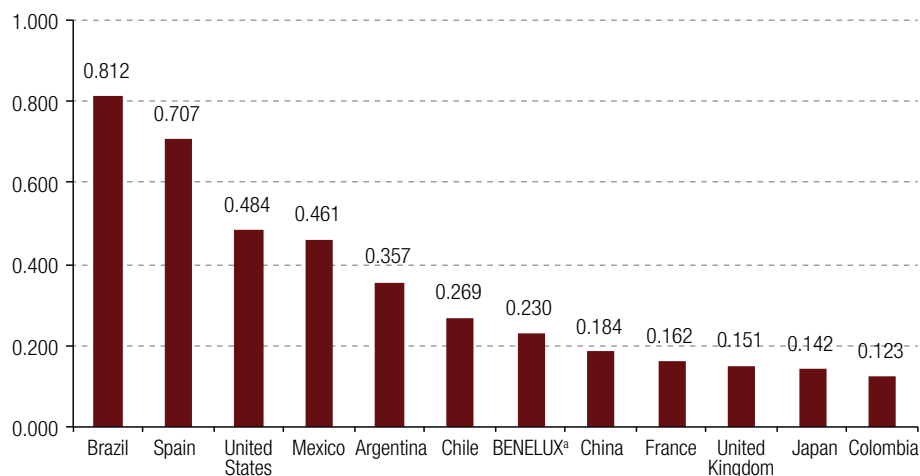
Source: Prepared by the authors.

ᵃ Economic union of Belgium, the Netherlands and Luxembourg.

Accordingly, despite the increase in the overall centralization of the network (largely owing to the role of Brazil), other centers of information exchange and dynamism in the transfer of experiences, learning, technology and other factors have emerged during this period, aside from the traditional players such as Spain, United States or Mexico. Economies as diverse as Colombia, Chile and China have started to gain relative importance in the interactions studied.

3. Mergers and acquisitions in the telecommunications, energy and finance sectors: regional dynamics and the role of Spain

When mergers and acquisitions are classified by the sectors of production in which the firms operate, a differentiated analysis reveals the behaviour patterns pertaining to each sector. In this case, the services sectors in which Spanish transnational companies have been most dynamic in Latin America were selected, namely telecommunications, energy and finance (see table 4 and figure 4).

Table 4
Latin America: structural characteristics of the networks of cross-border business mergers and acquisitions, selected sectors

	Energy	Finance	Telecommunications	Global
Density	0.104	0.108	0.066	0.279
Total degree centrality	0.036	0.040	0.032	0.081

Source: Prepared by the authors.

Figure 4
Latin America: networks of cross-border business mergers and acquisitions, selected sectors, 1999-2012

A. Energy

B. Finance

Ángeles Sánchez Díez, Pablo Galaso Reca and José Manuel García de la Cruz
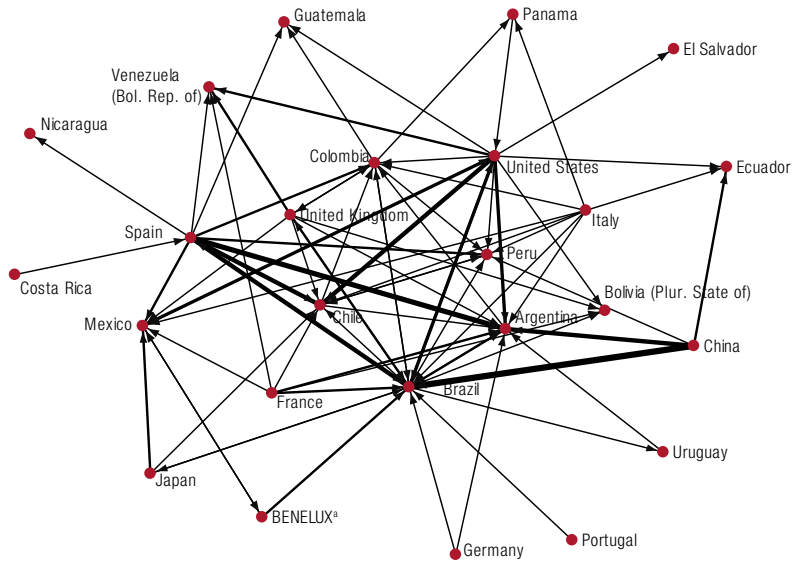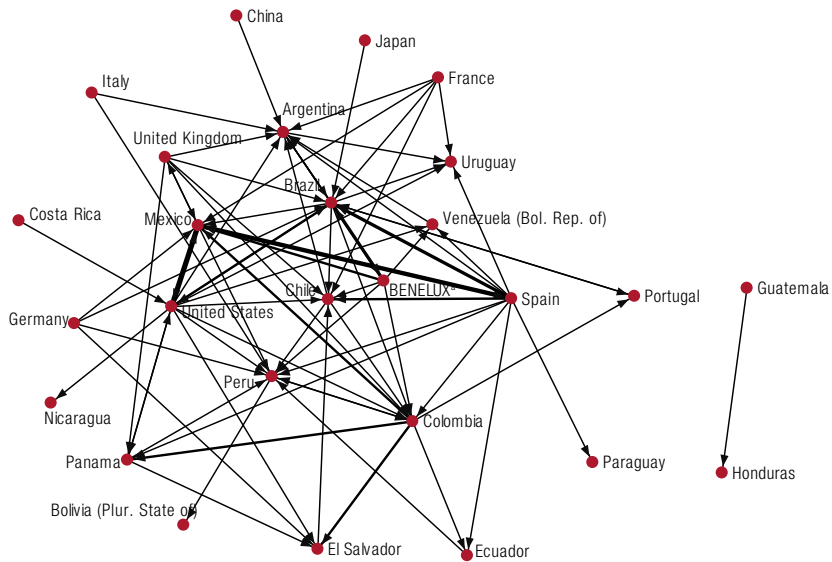
Figure 4 (concluded)

C. Telecomunications

Source: Prepared by the authors.

[a] Economic union of Belgium, the Netherlands and Luxembourg.

The analysis of the structural characteristics of these networks (see table 4) leads to the following conclusions:

- The interaction between countries, as measured by density, is greatest in the energy and finance sectors, and least in the telecommunications sector. Whereas many banking and energy firms have been set up in Latin America, the telecommunications market is basically controlled by two large operators.

- The sectoral networks display lower concentration levels, as measured through centralization, than those of the total network. This means that the mergers and acquisitions in the energy, telecommunications and finance sectors involve a larger number of active countries, since most countries have liberalized those public utility sectors to permit the entry of foreign-owned companies, which have set up businesses in those economies by purchasing existing assets.

In each of the sectors considered, it is possible to discern Spain's position relative to all countries, and with respect to those with which it disputes the most influential positions, particularly Brazil and the United States —although there are specific features in each sector.

In the energy sector, Spain has been more active than the United States and Brazil in acquiring Latin American firms, as measured by out-degree centrality. Companies such as Empresa Nacional de Electricidad, S.A. (ENDESA), now owned by the Italian firm ENEL, along with Iberdrola, Gas Natural and Unión Fenosa —the last two now merged as Gas Natural Fenosa— have expanded vigorously in the region. In the 1990s, these firms started to change their business structure to adapt to the legislative changes arising from the liberalization of the electricity and gas sectors imposed by the European Union. This meant not only operating in a climate of competition, but also vertically separating the generation, distribution and marketing segments. In addition, new technological developments associated with the generation of electricity through combined-cycle turbines linked the electricity and gas businesses. The final outcome was a process of mergers and acquisitions in the national markets, followed by their international expansion. At the same time, regulatory frameworks in the electric power sector in Latin America were altered, although with differences between countries (Maldonado and Palma, 2004). In this context, the Spanish firms acquired major assets in the region, competing with a number of United States enterprises such as Enron and AES Corporation, which had begun

their operations in the region a decade earlier. The energy crisis in California and the bankruptcy of Enron in 2001 cleared the path for Spanish firms and other European enterprises, while also opening up space for Latin American firms themselves. Nonetheless, the regulatory frameworks prevailing in Latin America have generated major uncertainties and regulatory risks (ECLAC, 2004), so some of the larger or faster-growing economies currently pose major energy-security challenges.

With regard to eigenvector centrality in the energy sector, Spain has been slightly overtaken by Brazil in the position of central country (node) although it is well ahead of the United States. As explained above, this indicator shows the connection with other countries (nodes) that are central in the network. It should be recalled that while Brazil privatized assets in the energy distribution and transmission segments, its privatization of the generating companies was interrupted by the energy crisis of 2001-2002 (ECLAC, 2012). In addition, several Brazilian public enterprises have crossed borders, such as Petróleo Brasileiro (Petrobras), which, along with Gas Natural Fenosa, is also one of the few companies in the region to show significant levels of integration between electric power generation and natural gas activities.

In the financial sector network, Spain stands out as an asset-purchasing country, according to the out-degree centrality measure, which puts it ahead of both the United States and Brazil. The Spanish banks have purchased numerous assets as they found a market in Latin America for the expansion of their activities.[7] Moreover, those investments enabled them to exploit their resources and organizational and technological capacities in a global way, thus diversifying risk (ECLAC, 2012). Entities such as Banco Santander and Banco Bilbao Vizcaya Argentaria (BBVA) are present in nearly all countries, having entered by taking over national banks. Although the penetration models of the two entities were different, there are several common features, such as the search for a potentially very broad market, since banking services are much less widely used in Latin American countries than in Spain and other countries of the European Union (Sánchez Díez, 2002). The lower in-degree centrality of Brazil can be explained by the importance of the public banking sector and local groups, which have made it difficult for foreign banks to enter the market. Regional competitors (both Colombian and Brazilian) have also appeared, however, along with European ones (British and Dutch), which are gaining substantial penetration in the region.

The analysis of eigenvector centrality in the financial sector shows that Spain has a more peripheral position than the United States, reflecting the historical presence of American banks in the region, which have maintained and consolidated their position over the years. Although foreign banks have a very large presence, representing 40% of the total (BIS, 2010), the international financial crisis has not seriously affected the region's banking sector (ECLAC, 2012, p. 144). Financial innovations and the deregulation processes themselves have been implemented cautiously, to avoid the critical experiences of the financial and banking debacles that have occurred in the last few decades. This is considered an intrinsic strength of the region's financial systems (Marshall, 2011).

In the telecommunications sector, Brazil and Spain are the network's most central countries, according to their total degree centrality, followed by Mexico, Portugal and Argentina. Brazil leads in terms of in-degree centrality, and Spain in terms of out-degree centrality. In the 1990s, telecommunication firms in Latin America were essentially public-sector monopolies with major infrastructure and service quality deficiencies, stemming from the lack of financing during the "lost decade." Capital account liberalization and legislative changes, in conjunction with privatization processes, enticed foreign firms into the region, from both Europe and the United States. Yet the new century surprised the sector with a major technological crisis that triggered a profound restructuring of the business, and the universalization of Internet use in civilian life. This transformed the importance of the services provided, such that fixed-line telephony has lost ground to mobile telephony and data transmission.

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[7]   ECLAC (2012, p. 130) contains details of the most important acquisitions in 1990-2011.

Some of the United States firms sold their assets, prioritizing investments in the national market, and the resulting opportunities were exploited by the Spanish operator Telefónica, currently Movistar, and Mexico's América Móvil. Since then, although other smaller operators exist, these two large firms have disputed the Latin American telecommunications market. Both are vertically integrated enterprises that have grown with the aim of being national leaders with a clear vocation towards internationalization. The Spanish firm initially focused on the southern cone and fixed-line telephony, before later targeting the whole region and all telecommunications markets; while the Mexican firm started to expand via the larger economies (Brazil) or those that were closest (a number of Central American countries), focusing on mobile telephony (ECLAC, 2008). Brazil has been the battleground for both firms, where Movistar received support from Portugal Telecom to create Vivo, the country's largest mobile telephone operator.

V.   Conclusions and final remarks

The analysis of network indicators leads to the following conclusions for each of the aims posed in the study:

(i)   With respect to aim 1, on the analysis of the network structure of mergers and acquisitions, the following are the key results in terms of interactions between economies and the concentration of operations in a group of countries:

•   The structure of the global network in 1999-2012 consists of a small group of central countries, while the others are on the periphery; in other words, some countries are highly interconnected through merger and acquisition flows, while other less interdependent countries are located on the fringe of the reorganization of productive asset ownership. An analysis of the trend during the period studied shows that the highest levels of interaction were attained during two periods, 2007-2009 and 2010-2011. This reflects the coexistence of asset acquisition strategies among firms that have traditionally been major investors in the region, such as those from Spain and the United States, and the emergence of certain Latin American countries as investors.

•   Network concentration levels, measured through centralization, are highest between 2007 and 2011, which reveals the increased importance of a small group of countries in those years, compared to a more equal distribution in earlier years.

    Accordingly, a central nucleus is forming of countries that participate more actively in the reorganization of Latin American assets, such that the growth of productive capital is consolidating a hard core of economies in which the majority of spillover effects and potential transformations of the production fabric are concentrated.

(ii)   As regards aim 2, concerning the study of Spain's position in the merger and acquisitions network, compared to that of its potential competitors, the findings are as follows:

•   An analysis of the position occupied by each country in the network reveals four types of country: those with high levels of total centrality, either as investor countries or as recipients, which clearly include Brazil, the United States and Mexico; (ii) those, such as Spain, that occupy a central position in the network owing to their major role as investors; (iii) those that have a central position as a result of asset sales, such as Argentina, albeit with a very clear trend of diminishing importance over the years, and (iv) those that are on the fringe of the network, scarcely connected with the central countries of the periphery and united by very weak links.

•   Spain has maintained central positions in the network, particularly in the periods 1999-2000 and 2003-2004. The emergence of new competitors has not pushed it to the periphery, although it

has meant a transformation of the network structure, owing to the larger number of countries in central positions.

- Countries that are able to compete with Spain for the central network positions include Brazil in particular, owing to its capacity to attract investment and its increasing strength as an investor country, especially since 2007. Firms from the United States have taken over Latin American enterprises throughout the period, but that country's in-degree centrality has also increased since 2007, which shows that trans-Latin firms are acquiring assets in that economy. Colombia and Chile also stand out, although to a lesser extent, as recipients of investment through the transfer of assets to foreign hands, while their firms have also been internationalizing in recent years, targeting neighbouring economies in particular. Lastly, China was very active in 2010-2011 as an investor country, with many of its firms seeking raw materials abroad to fuel national economic growth.

  Accordingly, the conclusion is that Spain is maintaining its position at the centre of the network, although with growing and powerful competitors, including Brazil and United States, and probably China in the near future.

(iii)   Lastly, the analysis of the sectoral networks structure (aim 3) leads to the following conclusions:

- Countries display greater interactions in the energy and finance sector networks than in telecommunications.

- The sectoral networks are less concentrated than the overall network. Nonetheless, whereas in the finance sector there are larger differences between the in-degree centrality index and out-degree centrality, in the cases of telecommunications and energy, the two concentrations are similar.

  Substantial differences therefore exist between the sectoral networks, with greater interactions in the energy and finance sectors. Power is less concentrated in the sectoral networks than in the global one.

To summarize, the network structure of business mergers and acquisitions in Latin America shows the existence of a central core of countries that account for the bulk of asset purchase operations and, consequently, have greater access to the production of knowledge spillovers, technology transfer or learning of new forms of business organization. This means that many countries remain on the periphery of the network. Spain continues to hold central positions, albeit with powerful competitors.

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The Chilean peso exchange-rate carry trade and turbulence

Paulo Cox and José Gabriel Carreño

Abstract

In this study we provide evidence regarding the relationship between the Chilean peso carry trade and currency crashes of the peso against other currencies. Using a rich dataset containing information from the local Chilean forward market, we show that speculation aimed at taking advantage of the recently large interest rate differentials between the peso and developed-country currencies has led to several episodes of abnormal turbulence, as measured by the exchange-rate distribution's skewness coefficient. In line with the interpretative framework linking turbulence to changes in the forward positions of speculators, we find that turbulence is higher in periods during which measures of global uncertainty have been particularly high.

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I. Introduction

Between 15 and 23 September 2011, the Chilean peso depreciated against the United States dollar by about 8.2% (see figure 1). The magnitude of this depreciation was several times greater than the average daily volatility of the exchange rate for these currencies between 2002 and 2012.[1] No events that would affect any fundamental factor that influences the price relationship between these currencies seems to have occurred that would trigger this large and abrupt adjustment. This change occurred during a period of great global uncertainty in the context of the debate over the United States' debt ceiling that raged during the months of August, September and October of 2011, which was magnified by the Federal Reserve's announcement confirming that it expected to see a highly uncertain scenario for the United States economy in the week of 20 September.[2]

Before the debate about the debt ceiling had sparked increased global uncertainty and the sharp depreciation, there had been a moderate but sustained appreciation of the Chilean peso throughout August 2011. This gradual but steady appreciation, followed by a sudden and large depreciation, gave rise to relatively large skewness or asymmetry coefficients in the distribution of the daily variations in the exchange rate. In fact, the skewness coefficient for the Chilean peso/United States dollar exchange rate was 1.89 during the third quarter of 2011, when this event occurred.[3] Simultaneously, and most importantly, sudden movements in the forward positions held by non-resident traders (FPNRs) in the local currency market —primarily based on non-deliverable forward contracts (NDFs)— towards buyer positions[4] had reversed their trend during the preceding weeks and had begun to move towards seller positions (see figure 1), while interest rate differentials between the Chilean peso and the United States dollar were at record highs.[5] Indeed, comparing 16 Friday to 23 Friday of the following week, FPNRs had exhibited an accumulated drop of about US$ 2.3 billion, with an average daily drop of about US$ 570 million. This decrease represents almost twice the standard deviation of daily changes in net FPNRs for 2011.[6]

As is documented in work by Brunnermeier, Nagel and Pedersen (2008), changes in these positions can be associated with investors that use forward contracts in the currency market for speculative purposes, a popular strategy known as "carry trade".[7] Currency carry trades such as the

[1]   The standard deviation of the daily variation in the exchange rate is 0.79 in the sample (2000-2012). The average daily depreciation during this episode was approximately 2.05%, which corresponds to more than 2.5 standard deviations.

[2]   The Economic Policy Uncertainty (EPU) Index for the United States built by Baker, Bloom and Davis (2015) reached a record high during August and September of 2011, substantially exceeding, for example, the level reached during the terrorist attack of 11 September 2001. For further details, see figure A1.1 in the annex or consult www.policyuncertainty.com/us_daily.html.

[3]   This figure represents the second-highest record in the sample (2002-2012). During the first quarter of 2011, the skewness of the daily variation in the exchange rate was 3.04. The monthly skewness coefficient for September 2011 was 1.02, the fourth-highest skewness coefficient in the entire sample.

[4]   This is from the point of view of local banks. A buyer carry-trade position, as seen from the point of view of a local bank, corresponds to a seller position from an investor's point of view.

[5]   The average interest rate differential was 5.1% in 2011 and 5.4% in the third quarter of that year. For the 2002-2012 sample as a whole, the average interest rate differential between the Chilean peso and the United States dollar was 2%. Figure A1.3 in the annex plots the relationship between the three-month interest rate differential and the level of exchange-rate volatility for the entire period.

[6]   Note that 2011 was a particularly volatile year in terms of FPNRs.

[7]   A carry trader is an investor who borrows in a low-interest-rate currency and invests in a high-interest-rate currency. Since this type of investment is not undertaken for hedging purposes, it is associated with speculation. Carry trade and greater foreign participation in local currency markets are recent phenomena that have exhibited significant growth (see Alfaro and Kanczuk (2013)). Of course, there are also mechanisms other than trades in the non-deliverable forward contract (NDF) market that can be used to conduct carry trade investments. (We thank José de Gregorio for drawing our attention to this point.) And, of course, not only foreign investors are implementing these strategies: domestic traders speculate as well. We believe, however, that FPNRs provide a better approximation of the correlations that we want to study here. In addition, the different types of investors are easier to identify, especially in the Chilean case, by examining trades in the NDF market.

ones that are of interest to us here are an investment strategy whereby speculators exploit the interest rate differentials between two currencies by taking short positions (debt) in a low-interest-rate currency (the "funding currency") in order to invest ("go long") in a higher-interest-rate currency (the "investment currency"). Because this strategy is not used to cover exchange-rate depreciation nor employed for hedging trading positions, the carry trade is associated with speculative behavior. As is the case with any form of speculation, the carry trade is a double-edged sword. On the one hand, it provides liquidity to the foreign-exchange market and potentially leads prices towards their fundamental level, thereby improving the performance of trade and financial activities. On the other hand, speculation in the exchange rate can create risks that would not be present in situations in which these strategies could not be applied. These risk scenarios and their implications for monetary policy are heavily influenced by the institutional framework and the economic factors that are specific to each economy.[8]

Figure 1
Chilean peso/United States dollar exchange rates and forward positions of non-resident traders in non-deliverable forward contracts, 2011
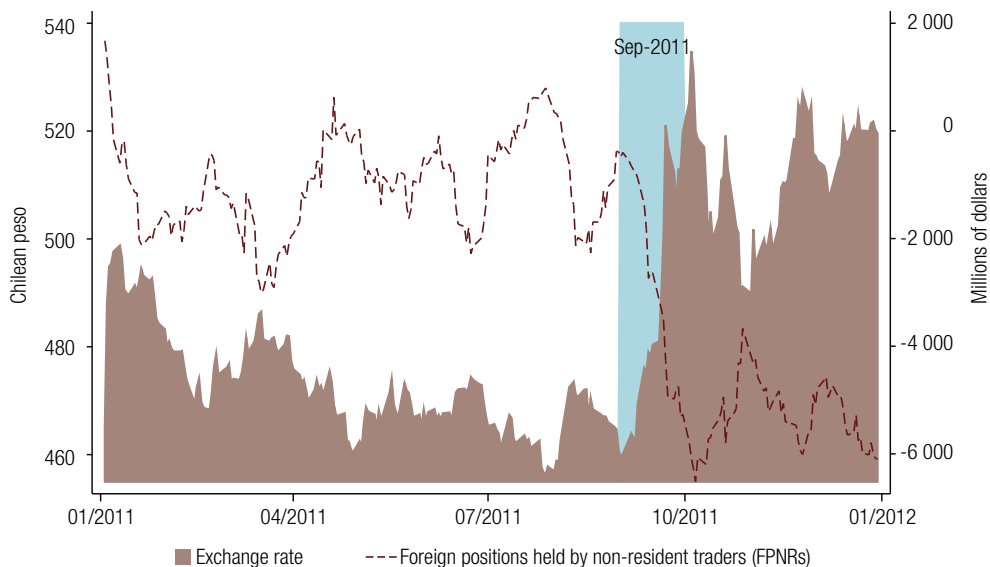
Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.

Like many other emerging economies, the Chilean economy is not free of these risks. To assess the potential risks (or lack of them) in the Chilean case, one of the many factors to be determined is whether the persistently large interest rate differentials existing between Chile and other advanced economies in recent years have given rise to a greater use of carry trade and, if so, whether this has sparked greater turbulence in the exchange rate or not.

Closely following work by Brunnermeier, Nagel and Pedersen (2008), the present study offers evidence on carry trade between the Chilean peso and the main currencies of advanced economies (particularly the United States dollar) that are traded in the local formal exchange market, and their effect on the skewness of daily changes in the exchange rate (our measure for turbulence). Based on a unique database of on-shore currency derivative transactions in the Chilean market, we study the behaviour of the forward positions of foreign investors and plot the relationships between those positions, on the

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[8] Recent developments in such strategies have attracted a great deal of attention in the aftermath of the global financial crisis of 2008 and have fuelled concerns about the effect that low interest rates in developed economies could have on currency stability and the effectiveness of monetary policy in emerging countries. For an exploration of the implications of this recent development for the conduct of monetary policy, see Plantin and Shin (2011).

one hand, and interest rates differentials and the skewness of daily changes in the peso/dollar and peso/euro exchange rates, on the other.[9] The main results of these comparisons are of substantial interest and have not been fully documented before now.

First, we provide evidence that the Chilean peso has been exposed to crash risk: positive interest rate differentials are correlated with a positive conditional skewness of changes in exchange-rate movements. Second, there is a correlation that suggests a causal relationship between movements in the forward net position of foreign investors and the asymmetry coefficient (see figure 1). Finally, we show that an increase in global risk or risk aversion[10] coincides with reductions in the net forward positions of foreign investors, as suggested by the interpretation presented in Brunnermeier, Nagel and Pedersen (2008) and Brunnermeier and Pedersen (2009).[11] Crash risk discourages speculators from taking positions that are large enough to reverse interest rate parity and move it towards its equilibrium, which in turn explains the forward premium puzzle.

The rest of this article is organized as follows. In section II, we briefly review the relationship between the forward premium puzzle and carry trade and look at how this relationship has been addressed in the literature. In section III, we describe the data and present preliminary evidence on the relationship between carry trade, FPNR movements and the exchange rate. Section IV presents our results and section V concludes.

II.   Related literature: the forward premium puzzle

One of the best-known empirical puzzles in the macroeconomic and financial literature is the forward premium puzzle,[12] which represents a violation of uncovered interest rate (UIP) parity. According to economic theory, when applied to the particular case of the foreign-exchange market, speculators in frictionless markets will arbitrage away any profit opportunity for exploiting interest rate differentials between two currencies. According to this principle, the currencies of economies with high interest rates (investment currencies) should tend to depreciate relative to the currencies of economies with lower interest rates (funding currencies). This hypothesis has been disproven empirically, however: on average, investment currencies tend to appreciate relative to funding currencies.[13]

In the seminal work by Engel (1996), which discusses this puzzle, the author concludes that traditional friction-free economic models are unable to solve this puzzle and suggests that new models which take into account other phenomena, such as the peso problem,[14] transaction costs or crash risk, among several other alternative hypotheses, should be seriously considered.

Thus far, several authors have attempted to explain the puzzle by following Engel's recommendation. Bacchetta and Van Wincoop (2010) and Mitchell, Pedersen and Pulvino (2007) point to market frictions that can block capital arbitrage. The first of these studies concluded that the excess return from carry trade is due to the infrequent review of investment decisions by investors, while the second study finds that financial constraints on speculators' investments create situations in

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[9] Our study thus complements studies based on data from off-shore operations. This is particularly the case with regard to the study carried out by Brunnermeier, Nagel and Pedersen (2008).

[10] We use the Volatility Index (VIX), the spread between the London Interbank Offered Rate (LIBOR) and the Overnight Indexed Swap (OIS) rate (known as LOIS) and the Economic Policy Uncertainty (EPU) Index built by Baker, Bloom and Davis (2015) as proxies for uncertainty (2013).

[11] According to this interpretation, the impact of the event of September 2011 that was described above is not an isolated one.

[12] For a review and discussion of this literature, see Hodrick (1987) and Engel (1996).

[13] The hypothesis is shown to be false by, for example, Bekaert and Hodrick (1992). See the discussion of this subject in the seminal work of Fama (1984) or, more recently, in Burnside and others (2010).

[14] That is, the effects of inference of low-probability events that are not observed in the sample.

which price gaps remain in place for long periods of time. Burnside and others (2010) argue that the positive average pay-off on an unhedged carry trade reflects its "peso event" risk.[15]

Following an alternative argument, some authors suggest that carry trade returns reflect a form of compensation for the crash risk present in these strategies (Gyntelberg and Remolona (2007), Lustig, Roussanov and Verdelhan (2008), Gromb and Vayanos (2010) and Jurek (2014)). For instance, Jurek (2014) finds that the crash risk premium accounts for at least one third of the excess return on currency carry trades. In line with this approach, Brunnermeier, Nagel and Pedersen (2008) study carry trades and currency crashes involving the United States dollar and apply the more general theoretical framework proposed in Brunnermeier and Pedersen (2009) in order to explain general liquidity problems. In this study, we test the hypothesis suggested in Brunnermeier, Nagel and Pedersen (2008) and in Brunnermeier and Pedersen (2009). This hypothesis says that sudden and abrupt exchange-rate depreciations which cannot be tied to news events concerning fundamentals are caused by the unwinding of carry trades when speculators approach the thresholds of their funding constraints. According to this hypothesis, large interest rate differentials encourage speculators to take positions which, in the absence of frictions, would bring profitable opportunities to an end. However, the crash risk discourages the same investors from taking positions that would completely close that profit window.

III. Data and preliminary evidence

In the following section, which deals with the results of our research, we will focus on the Chilean peso exchange rate *vis-à-vis* the most commonly traded foreign currencies in the Chilean local market: the United States dollar and the euro.[16] In this section, however, we will also look at other currencies for which data are available,[17] in addition to the dollar and the euro. We collect daily nominal exchange rates against the Chilean peso and three-month interest rates for the following currencies:[18] the United States dollar, the euro, the British pound, the Brazilian real and the Australian dollar. These are the most commonly traded currencies in the Chilean on-shore market.[19] We consider the period running from the first quarter of 2002 to January 2012.

1. The variables

We define the following variables:

Logarithm of the nominal exchange rate (s_t): the exchange rate corresponds to the number of Chilean pesos that are equivalent to one unit of foreign currency. The log of the nominal exchange rate is then defined as:

$$s_t = log(nominal\ exchange\ rate)$$

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[15] The term "peso event" refers to an event that is unlikely to occur but that would involve a large change if it were to happen.

[16] These two currencies comprise approximately 99% of all transactions involving foreign investors in the local market.

[17] This allows us to analyse whether currencies with significant interest rate differentials with respect to the Chilean peso are used as investment or funding currencies in carry trade investment strategies. According to the hypothesis advanced by Brunnermeier and Pedersen (2009), we should not observe their use for this purpose when the incentive for undertaking this type of investment strategy is low.

[18] To build the three-month interest rate differentials with respect to the Chilean peso, we use the 90-day prime rate cap and the 90-days British Bankers' Association (BBA) London Interbank Offered Rate (LIBOR) for the United States dollar, the British pound, the euro and the Australian dollar. Exchange rates are obtained from Bloomberg terminals.

[19] In an earlier version of this study, we also included the Peruvian sol, the Colombian peso, the New Zealand dollar, the Norwegian krone, the Mexican peso, the Japanese yen and the Canadian dollar. The results discussed in this section are consistent with the inclusion of these currencies.

Carry (interest rate differential) $\left(i_t^* - i_t\right)$: the difference between i_t^*, which denotes the logarithm of the domestic rate in period t (the foreign interest rate from the point of view of a foreign investor), and i_t^*, which denotes the logarithm of the foreign (investor's) economy's interest rate (associated with the currency being exchanged).

Carry trade return (z_t): The *ex post* return on investment denominated in a foreign currency that is financed with domestic debt (in pesos), where:

$$z_{t+1} = i_t^* - i_t - \Delta s_{t+1}$$

where:

$\Delta s_{t+1} = s_{t+1} - s_t$. The return is calculated from the perspective of a non-resident trader who invests in Chilean pesos and finances his/her positions in a foreign currency. Thus, the final return from this strategy is computed in units of foreign currency.

Asymmetry or skewness coefficient ($Skew_t$): the skewness coefficient (third moment of the distribution) of the daily change in the exchange rate, expressed on a quarterly basis. A negative skewness indicates that the tail on the left side of the probability distribution is longer than the one on the right side and that the bulk of the values lie to the right of the variable's mean (see figure 2). A positive skewness, conversely, indicates that the tail on the right-hand side of the probability distribution is longer than the tail on the left-hand side and that the bulk of the values lie on the left-hand side of the distribution. A zero value indicates that the values are equally distributed on both sides of the mean.

Figure 2
Distribution with negative skewness

Return (*in percentages*)
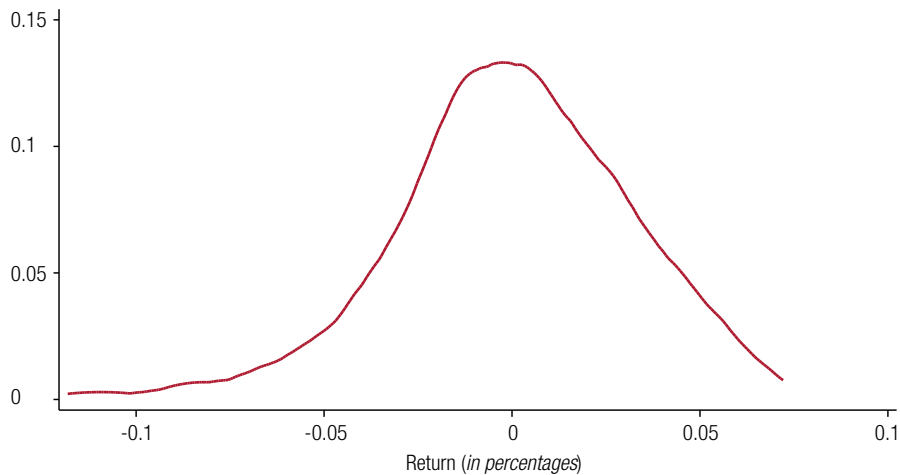
Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.

Foreign investors' non-deliverable forward net positions ($FPNR_t$): the net positions of non-residents in currency derivatives between the foreign currencies and the Chilean peso using NDFs.[20] For this variable, we have relevant information only for transactions based on the exchange rates between the Chilean peso and the United States dollar and between the Chilean peso and the euro.

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[20] In 2013, the Bank for International Settlements Triennial Central Bank Survey showed that NDFs constitute only a fifth of the global foreign-exchange market in outright forwards and a tiny fraction of overall foreign exchange trading. In the case of the Chilean peso, NDF transactions represent almost 90% of the transactions in the forward market (Salinas and Villena (2014)).

Throughout this study, FPNRs are analysed from the perspective of local banks. Accordingly, non-residents' buyer positions represent trades in which these investors are buying Chilean pesos or selling dollars (Chilean banks, on the other hand, are buying dollars or selling Chilean pesos). Positive values for this variable indicate that the foreign currency (dollar) is being used as a "funding currency" and the domestic currency (the peso) as an "investment currency."

This variable is the most important one in our study and is used as a proxy for carry trade activity. To construct this variable, we use information reported to the Central Bank of Chile on a daily basis[21] by commercial banks operating in the formal exchange market.[22] Chile requires all currency derivative transactions (mostly NDFs) to be reported to the central bank by the entities authorized to conduct such operations in the formal exchange market. Those reports must detail the counterparts' identities, the notional amounts, the type of compensation scheme, the expiration date and the price. Thus, the opacity that usually characterizes information in over-the-counter markets is not a factor here thanks to the nature of foreign exchange regulations in Chile.[23]

2. Carry trade and the asymmetry of changes in the exchange rate: preliminary evidence

Table 1 shows summary statistics for the main variables in our study for the five major currencies traded in the local currency derivative market. We note that there is a positive cross-section correlation between the average interest rate differential $i_t^* - i_t$ and the average excess return z_t, which implies a violation of the uncovered interest parity (UIP) condition. The largest average excess return (in United States dollars), for example, had the highest average interest rate differential in the sample.

An investor taking a long carry-trade position in Chilean pesos, financed with debt in United States dollars, would have earned the quarterly average of the interest rate differential —of 0.004 (i.e. an annualized return of 1.61%)— plus a quarterly excess return on the exchange rate of about 0.003 (1.21% per year) during the period covered in our sample. However, at the same time, the investor would have been exposed to a positive skewness of 0.049.[24]

In the last row of table 1, we report FPNRs. A negative position means that, in the aggregate, speculators hold a net selling position in dollars; conversely, positive positions imply that this group of investors constitute a net buyer. If the positions were predominately composed of carry trade, net FPNRs would be positive, on average. This is not what we observe in table 1. Thus, there are transactions other than speculative ones, such as hedging, for instance, which overshadow the carry trade transactions. However, as mentioned above, what interests us here is not the level of FPNRs, but the change in these positions from one period to another and the relationship between these changes, on the one hand, and the interest rate differentials and the skewness coefficient, on the other.

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[21] The central bank collects information on spot transactions and derivative contracts that are concluded by banking firms and other institutions belonging to the formal exchange market in both the local and off-shore markets. This information is collected in accordance with the Compendium of Foreign Exchange Regulations Manual. For more details about the depth, liquidity and size of the formal exchange market and its recent development, see Ahumada and Selaive (2007) and Salinas and Villena (2014).

[22] The positions at the end of every month are published by the Central Bank of Chile.

[23] Throughout most of the period considered in our study, aggregate FPNRs were negative, which indicates that, in aggregate terms, the FPNR variable does not reflect the implementation of carry trade strategies. However, it is the *change* in these positions which interests us here.

[24] We also note from the table that the bulk of "speculative" (that is, involving foreign investors) transactions are between the Chilean peso and the United States dollar, with a smaller volume of transactions involving the Chilean peso and the euro. Other currencies will not be discussed in the remaining sections of this study because the transactions involving these currencies are so sporadic.

Table 1
Descriptive statistics (means): quarterly data for the period 2002Q1-2012Q1

	USD	EUR	GBP	BRL	AUD
Δs_t	-0.003	0.001	-0.003	0.001	0.005
z_t	0.007	0.002	0.004	-0.01	-0.006
$i_{t-1}^* - i_{t-1}$	0.004	0.002	0.001	-0.01	-0.001
Skewness	0.0049	0.113	0.076	-0.023	-0.081
FPNR	-2 290	-49	…	…	…

Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.
Note: USD: United States dollars; EUR: Euros; GBP: British pounds; BRL: Brazilian reais; AUD: Australian dollar. Δs_t is the change in the logarithm of the exchange rate (Chilean pesos per unit of foreign currency). z_t, when the three-month interest rate differential is positive, is the return on investment in a long position in the local currency financed by a loan in foreign currency. When the difference is negative, it is the reverse. FPNRs are the net (long-short) forward positions of non-residents, in billions of pesos, considering NDF contracts only; FPNR data are for United States dollars since 2003Q1. Data for euros are for the period beginning in 2006Q2. A positive net FPNR implies that non-residents are, in the aggregate, making commitments to buy Chilean pesos in order to engage in carry trade.

Using information from table 1, in figure 3 we show the relationship between the carry trade, its return (panel 3A) and the skewness coefficient (panel 3B).[25] In line with the uncovered interest parity condition, the average return must be zero. However, there is a positive correlation between the average interest rate differentials and the excess return, which violates that condition (panel 3A). This relationship is in line with work done by Jurek (2014), who, using a sample of the Group of Ten (G10) currencies, finds that currency carry trade delivers significant excess returns, with annualized Sharpe ratios equal to or greater than those of equity markets (1900-2012). The evidence shown in both table 1 and figure 3 (panel 3B) also suggests that there is a positive relationship between skewness and average interest rate differentials. This positive correlation implies that the carry trade is subject to a positive skewness (crash risk) i.e. the Chilean peso is exposed to a sharp depreciation against the United States dollar and other currencies. We also note in table 1 that the skewness coefficient is negative for those currencies for which the interest rate differential is negative, such as the Brazilian real and the Australian dollar. This is consistent with the interpretation presented by Brunnermeier and Pedersen (2009).

Another way of obtaining evidence on excess returns and their correlation with the asymmetry coefficient is by analysing the distribution of the excess return z_t conditional on the interest rate differentials $i_{t-1}^* - i_{t-1}$ (figure 4), with the observations grouped according to the following ranges of differentials: $i_{t-1}^* - i_{t-1} \leq$ -0.0035; -0.0035 $\leq i_{t-1}^* - i_{t-1} \leq$ -0.0035; and $i_{t-1}^* - i_{t-1} \geq$ -0.0035.

We note from figure 4 that, in the case of significant interest rate differentials (blue dashed line),[26] the distribution of the excess return is, on average, positive and exhibits a long tail on the left-hand side, evincing the asymmetry of the change in the exchange rate (skewness). Conversely, when interest rate differentials are negative, we do not observe skewness in the return's distribution, since, for these currencies, the incentive to conduct carry trade with the Chilean peso is weaker, as is the case with the Brazilian real, as well. With respect to interest rate differentials around zero, we see that the distribution is centered at zero and symmetrical. All this confirms the existence of crash risk in carry trade transactions.

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[25] For more details on all the values observed, by currency and quarter, see figures A1.1 and A1.2 in the annex.

[26] Both the United States dollar and the euro belong to this group.

Figure 3
Cross-section empirical return, skewness and three-month interest rate differentials,
quarterly data for the period 2002Q1-2012Q1
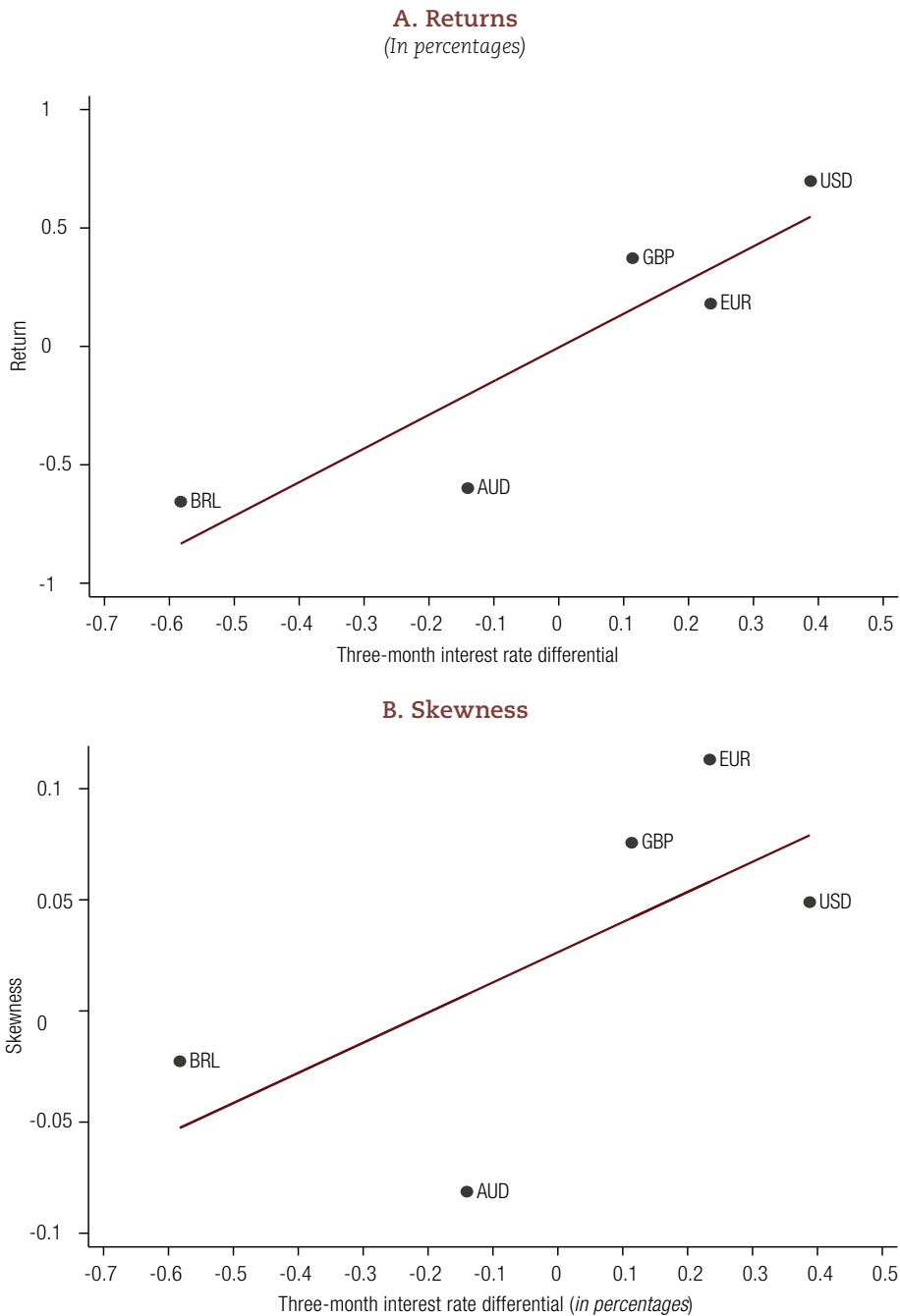
A. Returns
(In percentages)

B. Skewness

Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.
Note: USD: United States dollars; EUR: Euros; GBP: British pounds; BRL: Brazilian reais; AUD: Australian dollar.

Kernel distribution of the return z_t as a function of interest rate differentials
(after removing fixed effects) for the period 2002Q1-2012Q1
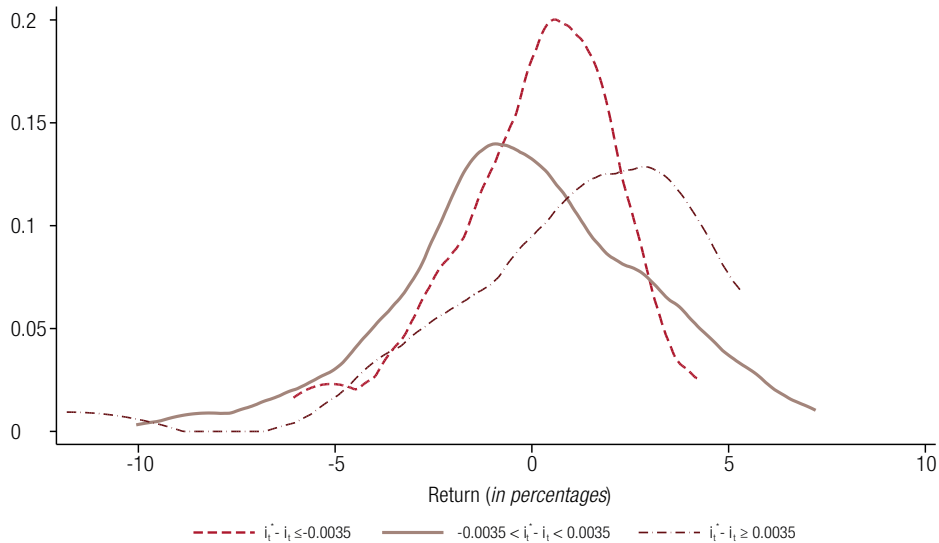
Return (*in percentages*)

- - - - $i_t^* - i_t \leq$ -0.0035   ———— -0.0035 < $i_t^* - i_t$ < 0.0035   – - – · – $i_t^* - i_t \geq$ 0.0035

Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.
Note: Data are for the United States dollar, euro, British pound, Australian dollar, Brazilian real, Peruvian sol, Colombian peso, New Zealand dollar, Norwegian krone, Mexican peso, Japanese yen and Canadian dollar.

To sum up, the evidence provided here, for a cross-section of currencies, points to a positive relationship between interest rate differentials and the risk of currency crashes. However, these results correspond to correlations between variables that may or may not be systematically related. The following analysis will exploit the time series, adding information on speculators' transactions, with the object of testing empirically the hypothesis put forward by Brunnermeier and Pedersen (2009).

IV. Results

1. The relationship between the currency carry trade and turbulence in the exchange rate

We begin by studying the predictors of large corrections in the exchange rate. In particular, we focus on interest rate differentials and carry trade activity. As a proxy for crash risk, we use the skewness coefficient ($Skew_t$), the dependent variable. Using simple linear regressions, we test whether the FPNR coefficient and lags of the dependent variable are significant predictors of crash risk in relation to the exchange rate. An increase in interest rate differentials affects crash risk positively. We thus expect a positive sign for its coefficient. A buyer FPNR position (carry trade activity with the Chilean peso), moreover, increases this risk, and we therefore expect a positive sign in this coefficient as well. Finally, we expect a positive sign in the lagged dependent variable's coefficient, as a higher level of risk in the past should discourage very aggressive short positions in the present.

Specifically, we consider the following panel regressions:

$$Skew_{jt+1} = \beta_1 Skew_{jt} + \beta_2 \left(i_{jt}^* - i_{jt} \right) + \beta_3 FPNR_{jt} + \beta_4 X_{jt} + \alpha_j + \varepsilon_{jt} \tag{1}$$

where j is a given country and t a given quarter. X_{jt} is a vector of controls that, depending on the specification, may include: z_{jt}, the carry trade's return; $BCCH_t$, which is a dummy that takes a value of 1 if, in the corresponding quarter, the central bank either announced a currency intervention or intervened in the exchange market, and 0 if not; $log(Copper)$, the log of the price of copper (cents of United States dollars/pound) reported in the London Metal Stock Exchange; $log(Oil)$ the log of the nominal price of oil reported in the New York Mercantile Exchange (NYMEX); the Emerging Markets Bond Index (EMBI), which is the weighted average of sovereign spreads for a large group of emerging-market economies and is intended to control for country risk changes; and, finally, α_j and ε_{jt}, which correspond to the currency's fixed effect and the disturbance term, respectively. The results are shown in table 2.[27] The first three columns show that the carry trade $(i_t^* - i_t)$ is a strong predictor of future skewness. We also find that the lag of the $Skew$ variable has a negative coefficient, which suggests a reversion to the mean in this variable. We also find that the FPNR coefficient is positively correlated with future skewness.

Table 2
Currency crash predictors

	$Skew_{t+1}$	$Skew_{t+1}$	$Skew_{t+1}$	$Risk\ Reversal_t$
$i_{t-1}^* - i_{t-1}$	57.34***	57.07***	53.31***	113.9***
	(13.22)	(12.82)	(0.501)	(9.182)
$FPNR_t$	0.0373***	0.0304***	0.128***	-0.257***
	(0.00153)	(0.00924)	(0.0443)	(0.0148)
$Skew_t$	-0.148***	-0.151***	-0.186***	-0.461***
	(0.0248)	(0.0154)	(0.0415)	(0.0688)
$BCCH_t$	0.848***	0.831***	0.765***	-0.618***
	(0.0609)	(0.0369)	(0.0745)	(0.0334)
z_t		2.651	1.631	-14.57***
		(3.016)	(1.092)	(2.545)
$log(Copper)$			0.338***	-0.560
			(0.119)	(0.836)
$log(Oil)$			-0.108	1.842***
			(0.119)	(0.448)
EMBI			-0.176	1.258***
			(0.427)	(0.130)
Observations	62	62	62	51
R^2	0.240	0.254	0.304	0.531

Source: Prepared by the authors, on the basis of information from the Central Bank of Chile and data from Bloomberg.
Note: Panel regressions with country fixed effects and quarterly data for 2003Q1-2012Q1. FPNR data include NDF contracts only, in billions of dollars. Risk reversals are the implied volatility difference between one-month foreign currency call and put options. Panel data are for United States dollars starting from the first quarter of 2002 and for euros starting from the second quarter of 2006. Risk reversal data are for United States dollars starting from the first quarter of 2005 and for euros starting from the third quarter of 2006. Clustered standard errors are shown in parentheses: *** $p<0.01$; ** $p<0.05$; * $p<0.1$.

We have included a dummy variable in all specifications for the central bank's intervention, which is positive and significant. Announcements or interventions by the central bank have been shown to increase the risk of currency collapse. To the extent that increases in $Skew_{t+1}$ arising from

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[27] To facilitate the reading and interpretation of the FPNR coefficients in this table, we have expressed this variable in billions of dollars. In the rest of the article, this variable is expressed in millions of dollars.

these announcements result in a lower degree of speculation, this result is useful for the control of exchange-rate turbulence, but must be analysed further.

We emphasize that all these results are consistent with evidence discussed in the previous sections. The second column shows that the currency return, z_t, has the expected sign (positive carry trade returns prompt investors to take positions), although the coefficient is not significant, possibly because the FPNR coefficient is strongly correlated with the return.

Note that the relationship between the FPNR coefficient and the skewness coefficient (Skew) is robust to the inclusion of fundamentals that are traditionally included in long-term models for the Chilean peso/United States dollar exchange rate, such as the price of copper, the price of oil and EMBI. Of these three variables, the price of copper is the only control variable that is statistically significant when regressed on the exchange rate's skewness.[28]

As another robustness check, we include the risk reversal variable (in the last column in table 2) as an alternative dependent variable for crash risk. The risk reversal is a measure of the implied volatility in the difference between a one-month foreign currency call option and a one-month foreign currency put option.[29] The price, then, reflects the cost of insurance coverage for changes in the exchange rate. The higher the crash risk, the higher its price will be. As expected, both the carry and the lagged skewness variable are significant and have the expected signs. The FPNR coefficient is negative and significant. It thus has the opposite sign to the theoretical one, possibly because the effect of a higher carry on this variable at time period t is already captured by the interest rate differential. These results are similar to those reported by Hutchison and Sushko (2013), who find a close link between risk reversals and speculative futures positions in the Japanese yen. Something similar occurs with return z_t.

Having established the empirical relationships among the carry trade, the skewness and FPNRs, in the following section we study the dynamic relationship among these three variables in detail using an autoregressive vector analysis.

2. Autoregressive vector analysis: the case of the Chilean peso-United States dollar carry trade

In this section we estimate a second-order autoregressive vector analysis focusing exclusively on transactions involving the United States dollar. This analysis allows us to study the systematic and dynamic relationship between the interest rate differential, the carry trade return, the positions of speculators (FPNR) and the skewness coefficient in the context of the carry trade between the Chilean peso and the United States dollar. We use monthly data for the period 2002m1-2012m1. The shocks underlying the impulse response function are based on a Choleski decomposition based on the following order: $i^*_t - i_t$, z, *Skewness* and FPNR.[30] All variables are filtered using the Hodrick-Prescott method for decomposition of the cyclical component of the series.

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[28] The fact that the Chilean peso has the status of a commodity currency suggests that the trend of this fundamental must be taken into account in the episode described in the introduction. Indeed, the price of copper, which is the main factor behind trends in the Chilean peso/United States dollar exchange rate, both in the short and long run, depreciated by about 12% during this episode (see the introduction). However, it should not be forgotten that this price is also affected by global uncertainty. It seems implausible, in fact, that uncertainty in the United States was caused by news concerning the price of copper during this episode. Evidence regarding the role of copper in recent misalignments of the exchange rate can be found in Wu (2013), where the author identifies several recent episodes in which fundamentals, in particular the price of copper, fail to explain distortions in short-term changes in the exchange rate. The evidence reported in this study backs up those findings, since it indicates that the identified exchange-rate misalignments have been caused by turbulence induced by the United States dollar/Chilean peso carry trade.

[29] The table shows a smaller number of observations for the specification using risk reversal as the dependent variable. This is due to data availability.

[30] This order is consistent with the theory on which our main workhorse hypothesis is based, i.e., that of Brunnermeier and Pedersen (2009).

Figure 5 reports the impulse response of a one-standard-deviation shock on the interest rate differential, using 95% confidence intervals. Panel 5C in figure 5 shows that the FPNR coefficient is closely related to the interest rate differential (panel 5A). Large spreads induce higher carry trade activity, triggered by an increase in the expected return for investors who use the Chilean peso as an investment currency and the United States dollar as a funding currency.

An increase in carry trade transactions has two effects on skewness ($Skew_t$). On the one hand, it generates an appreciation of the exchange rate; on the other, this appreciation incubates crash risk. The latter is reflected in the increased likelihood of abrupt adjustments in the exchange rate (an increase in positive skewness). Indeed, the bulk of the distribution of changes in the exchange rate shifts towards appreciation events, while the tail on the right-hand side of the distribution thickens due to the increase in crash risk events having a lower frequency but larger magnitude. This larger skewness effect is caused by a withdrawal (unwinding) of speculative positions, which in turn induces a "spiral loss." This increase in the risk of collapse leads investors to unwind positions or to not continue investing as much as they did before the shock. Panel 5D in the figure shows that, as the volume of carry trade transactions grows bigger (panel 5C), we first observe an increase in the skewness coefficient, followed by a gradual decline which converges towards pre-shock levels as the positions of the speculators recoil towards their initial level.

When interpreting the results of the autoregressive vector analysis, it is important to bear in mind that the impact of carry trade on the exchange rate would probably be much larger if we conditioned it on periods that exhibited greater carry trade activity (periods, for instance, during which interest rate differentials were particularly large). Indeed, for much of the relevant period, i.e. between 2002 and mid-2007, currency derivatives were not being used to any significant degree, probably because they were not, at that time, an attractive investment. Throughout the period, this effect tends to reduce the impact captured by an unconditional autoregressive vector analysis on the sample as a whole.

Figure 5
Impulse response functions from an autogressive vector analysis (2) for a one-standard-deviation shock to interest rate differentials for the period 2002m1-2012m1

A. Impulse

Figure 5 (continued)

B. Return on carry trade

Months

C. Forward positions held by non-residents

Months
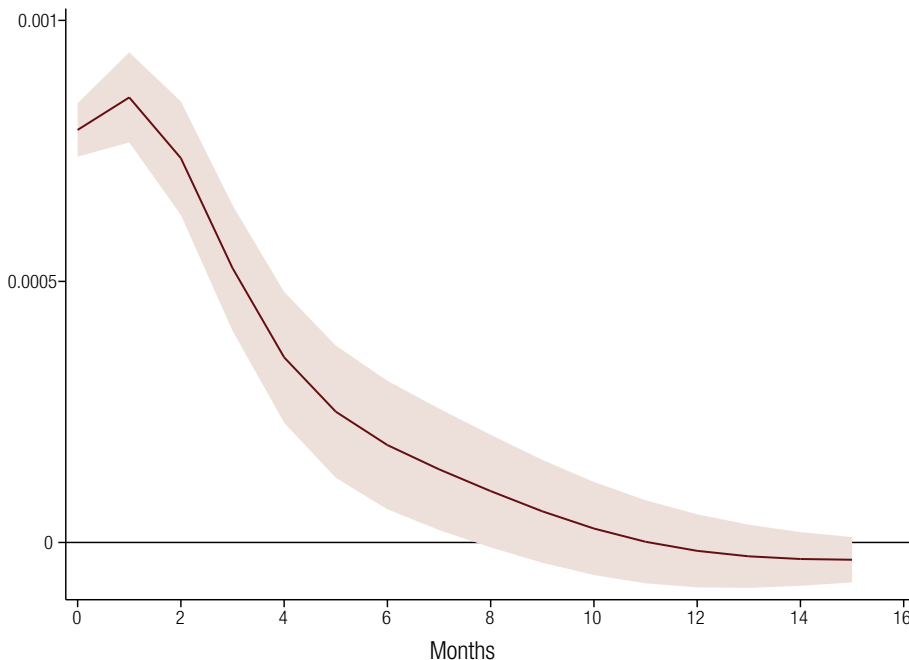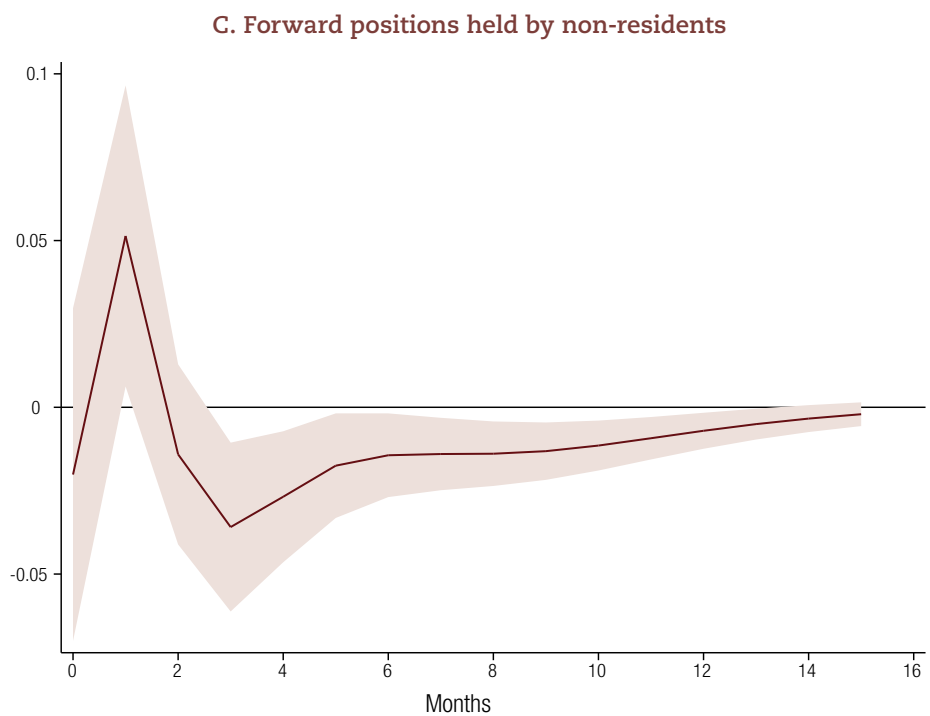
Figure 5 (concluded)

D. Skewness
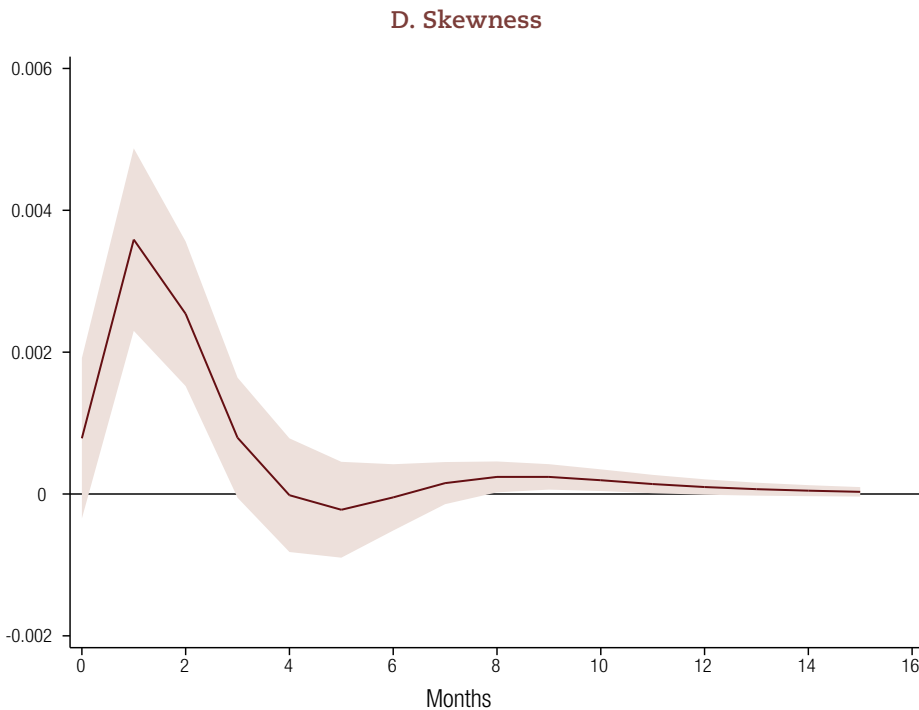
Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.

3. Liquidity risk and the unwinding of carry trades

In this section, we will take a closer look at speculators' behaviour. Specifically, we will study the factors that lead speculators to unwind their positions, such as sudden losses in their positions, an increase in funding margins or a reduction in their tolerance to risk. To this end, we use two measures for aggregate uncertainty that are normally associated with funding illiquidity and investors' risk aversion: (i) the volatility index of the Chicago Board Options Exchange (CBOE VIX), and (ii) the Economic Policy Uncertainty (EPU) Index built by Baker, Bloom and Davis (2015). Both measures of uncertainty are used as observable proxies that should be correlated with the factors mentioned above (see, for example, Anzuini and Fornari (2011) for the macroeconomic determinants of carry trade activity). We report results using the EPU Index only.[31] We consider the following panel regressions:

$$\Delta PFNR_{jt} = \beta_1 \Delta EPU_t \times sign\left(i^*_{jt-1} - i_{t-1}\right) + \beta_2 EPU_t + \beta_3 FPNR_{jt-1} + \beta_4 BCCH_t + \alpha_j + \varepsilon_{jt} \quad (2)$$

and

$$\Delta z_{jt} = \beta_1 \Delta EPU_t \times sign\left(i^*_{jt-1} - i_{t-1}\right) + \beta_2 EPU_t + \beta_3 BCCH_t + \alpha_j + \varepsilon_{jt} \quad (3)$$

where j is the country, t is time, α_j represents the country's (or currency) fixed effect, and ε_{jt} is the disturbance term. The expected sign for the coefficient of the interaction variable obtained from multiplying EPU and carry trade, under the hypothesis that the unwinding of positions is taking place and affecting crash risk, is negative in both regressions. Until now, we have ignored the direction of carry trade, given that interest rate differentials and other variables change sign when the direction of

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[31] Results using VIX are reported in the annex in table A1.1 and are consistent with those discussed here.

carry trade changes. Certainly this is not the case with VIX or EPU, and therefore we interact these two variables with the sign of interest rate differentials,[32] $sign(i^*_{t-1} - i_{t-1})$.

We estimate specifications in equations (2) and (3) using the fixed effects estimator. Table 3 presents the results for the EPU Index. The first two columns show that the impact of increased global uncertainty on contemporary changes in FPNRs (same month) is positive. This is reversed in the following month, as is to be expected in a context where an increase in global uncertainty causes an unwinding of carry trade positions. This implies that greater global uncertainty scenarios can lead to crash risk, but with a lag. The fact that the contemporaneous effect upon FPNRs is positive (column 1 in table 3) is not consistent with the hypothesis of Brunnermeier and Pedersen (2009). Our explanation for this seemingly puzzling result is the following. Shocks to uncertainty lead non-speculating investors to take positions in order to hedge their fundamentals-driven investments against higher volatility in the exchange rate. Readers will recall that FPNRs are predominantly used by these investors, who buy dollars. Subsequently, as long as persistent uncertainty shocks are narrowing funding margins in global financial markets, the changes observed in FPNRs will increasingly be accounted for by speculators who are unwinding their positions. This explains both the contemporaneous positive effect and the subsequent negative effect.

We also note from the table that increases in $FPNR_{t-1}$ predict a smaller change in $FPNR_t$, which could be explained by the presence of a larger crash risk.

Table 3
Monthly sensitivity of positions and carry trade returns to changes in the economic policy uncertainty index

	$\Delta FPNR_t$	$\Delta FPNR_{t+1}$	Z_t	Z_{t+1}
$\Delta EPU_t \times sign(i^*_{t-1} - i_{t-1})$	0.987***	-1.325***	5.25e-05***	-0.000170***
	(0.0433)	(0.153)	(1.74e-05)	(6.43e-05)
$\Delta FPNR_{t-1}$	-0.207***	-0.0686***		
	(0.00250)	(0.00494)		
$\Delta log(Copper)_{t-1}$	0.445*	-0.809***		
	(0.244)	(0.229)		
$BCCH_t$	-25.00***	47.55***	-0.00364***	-0.000604
	(5.678)	(15.00)	(8.36e-05)	(0.00311)
Observations	179	175	183	179
R^2	0.041	0.014	0.007	0.033

Source: Prepared by the authors, on the basis of information from the Central Bank of Chile and data from Bloomberg and Economic Policy Uncertain Index [online] www.policyuncertain.com.

Note: Panel with country fixed effects and monthly data, 2002M3-2012M1 for United States dollars and 2006M8-2012M3 for euros. FPNR data include NDF contracts only. Monthly interest rate at the beginning of period t. Clustered standard errors are shown in parentheses: *** p<0.01; ** p<0.05; * p<0.1.

In the third and fourth columns, we indicate the relationship between uncertainty and expected returns. Global uncertainty is positively correlated with the return on investment in the second month following a change in uncertainty, which discourages carry trade. Considering all these results, taken together, we can draw the following conclusion: when risk tolerance declines, traders unwind their positions, inducing a reduction in FPNRs and in return z_t. The inclusion of the $BCCH_t$ dummy shows it to be an important explanatory factor for both the return on carry trade and changes in the FPNR variable, although its impact is short-lived.

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[32] Of course, this is only meaningful when the carry sign changes. Although the frequency of the occurrence of this event is low in the sample, there are quarters for which the direction of the carry between the Chilean peso and the United States dollar and between the peso and the euro actually does change from positive to negative and vice versa.

The inclusion of a fundamental variable, such as the price of copper, is relevant, but is not robust to the inclusion of other controls.[33] Given the strong contemporary relationship between the EPU Index and the carry trade excess return, we analyse to what extent indices for global uncertainty are able to predict future returns and to approximate changes in risk premiums (the foundation for the carry trade return). We consider the following econometric panel specification:

$$z_{jt+1} = \beta_1 \left(i_{jt}^* - i_t \right) + \beta_2 \left(\Theta_t \times sign \left(i_{jt-1}^* - i_{t-1} \right) \right) + \alpha_j + \varepsilon_{jt} \tag{3}$$

where j is the country, t time, αj represents the country's (or currency's) fixed effects and ε_{jt} is the disturbance term. Θ_t is the indicator of global uncertainty or liquidity. In addition to the global uncertainty measures (the Economic Policy Uncertainty (EPU) Index and the volatility index (VIX)), we consider the measure of liquidity provided by the spread between the London Interbank Offered Rate (LIBOR) and the Overnight Indexed Swap (OIS) rate (known as LOIS). Results from fixed effects estimations are shown in table 4.

Table 4
Exchange-rate return z_t regressed on carry trade $i_t^* - i_t$ and its interaction with the volatility index, the spread between the London Interbank Offered Rate and the Overnight Indexed Swap rate, and the economic policy uncertainty index

Excess return on	z_{t+1}	z_{t+2}	z_{t+3}	z_{t+1}	z_{t+2}	z_{t+3}	z_{t+1}	z_{t+2}	z_{t+3}
$i_t^* - i_t$	1.073***	0.924***	0.724***	1.077***	0.882***	0.694***	1.121***	0.974***	0.721***
	(0.297)	(0.244)	(0.122)	(0.293)	(0.258)	(0.110)	(0.275)	(0.209)	(0.141)
$\Delta VIX_t \times sign\,(i_t^* - i_t)$	-0.000991	-0.000326***	-0.000398						
	(0.000753)	(0.000124)	(0.000376)						
$\Delta LOIS_t \times sign\,(i_t^* - i_t)$				-0.0147**	-0.0222***	-0.0100***			
				(0.00739)	(0.00798)	(0.00140)			
$\Delta EPU_t \times sign\,(i_t^* - i_t)$							-0.000116***	-0.000154	-1.15e-05
							(2.23e-05)	(0.000163)	(9.92e-05)
Observations	236	234	232	236	234	232	236	234	232
R^2	0.147	0.052	0.042	0.091	0.116	0.043	0.073	0.066	0.028

Source: Prepared by the authors, on the basis of information from the Central Bank of Chile and data from Bloomberg and Economic Policy Uncertain Index [online] www.policyuncertain.com.

Note: Panel regressions with country fixed effects and monthly data are for 2002M3-2012M1. The data are for United States dollars and euros. The Volatility Index (VIX) measures "appetite" for risk. The spread between LIBOR and the Overnight Indexed Swap (OIS) rate (LOIS) is a measure for international liquidity, while the Economic Policy Uncertainty (EPU) Index is a measure of economic and political uncertainty. VIX, LOIS and EPU correspond to the final observation for each quarter. Clustered standard errors are shown in parentheses: *** p<0.01; ** p<0.05; * p<0.1.

One notable result is that the effect of uncertainty on returns is confirmed, and this is robust across the different measures for global uncertainty.[34] The effect is also shown to be persistent over time. An increase in uncertainty in   negatively impacts the return in the same quarter and in the following two quarters, although the impact eventually becomes non-significant over time. Overall uncertainty, therefore, is an important predictor of returns, while these effects are relevant in the short term. Evidence is consistent, then, with the hypothesis of Brunnermeier and Pedersen (2009).

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[33] For example, when including the lag of carry trade $(i_{t-1}^* - i_{t-1})$ in the regression, the price of copper is not significant, while the other estimated coefficients keep their signs and significance.

[34] Table A1.1 in the annex shows the results using VIX instead of EPU as an indicator for uncertainty.

4.   Where is carry trade activity more concentrated: in short- or long-duration contracts?

One of the advantages of the data sources used in this study is that they allow us to obtain detailed information about NDF contracts that is usually not available in other economies or markets. In particular, we are able to classify trading transactions (the notional amounts of NDFs) according to their duration. In this section, we use this information to identify the duration of the contracts for which the carry trade is most active in the NDF exchange market.

Table 5 depicts the differing carry trade effect on NDF transactions subscribed by non-resident agents according to their duration. A carry trade investment strategy should entail the adoption of longer positions by speculators following an increase in interest rate differentials. As shown in table 5, for short-term (<1 month) operations, the relationship between the carry trade and FPNRs (the dependent variable) is not statistically significant. In the medium run (i.e. for contracts ranging from one month to one year), however, we note that carry trade is significantly correlated to FPNRs, which is consistent with the evidence and analysis presented here and elsewhere (Ichiue and Koyama (2011)). We also notice that the sign accompanying the carry trade coefficient is reversed for longer durations, confirming that speculative activity is concentrated in medium-duration contracts.

Table 5
Forward positions held by non-resident traders, by term of the underlying non-deliverable forward contracts, regressed on carry $i_{t-1}^* - i_{t-1}$

	Forward positions held by non-residents, disaggregated by the term of the contract				
	7 days	8-30 days	31-180 days	181-365 days	>365 days
$i_{t-1}^* - i_{t-1}$	1.837	-2.381	90.495**	72.365***	-50.315*
	(1.495)	(10.656)	(44.102)	(24.934)	(27.207)
Observations	333	119	40	40	40
R^2	0.005	0.0006	0.11	0.21	0.04

Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.
Note: Ordinary least squares (OLS) regressions are for United States dollars for the period 2002M1-2012M1. Data for seven-day contracts are available from 2005 on. Forward positions of non-residents are expressed in billions of United States dollars. Interest rate differentials are for 7, 30, 180 and 360 days, respectively (weekly, monthly and quarterly data), with the latter being derived from 180-day rates. Interest rate differentials shown in the first and second columns are for the beginning of the week/month. In third and fifth columns, the interest rate differentials correspond to the beginning of the quarter. For contracts of more than one year in duration, the interest rate differentials are based on one-year interest rates. For Chilean pesos, the prime cap rate is used, while for United States dollars, LIBOR is used. Robust standard errors are shown in parentheses: *** p<0.01; ** p<0.05, * p<0.1.

V.   Conclusions

In this paper we have documented evidence on the use by foreign investors of currency carry trade investment strategies involving the Chilean peso and other developed-economy currencies; the use of this strategy has been encouraged by the large interest rate differentials between these currencies observed in recent years. This has had a significant impact, especially during periods of great global uncertainty, on the nominal exchange rate, which has gone through numerous episodes of turbulence. By using several econometric specifications, we show that a significant and robust explanatory factor in the distribution of daily changes in the exchange rate is the forward position of non-resident traders in the local market. Consistent with the interpretation of the relationship between the distribution's asymmetry coefficient and the futures positions of such investors, we also find that the risk of a currency collapse (crash risk) discourages investors from taking positions that would lead to the

restoration of parity rates. We also have shown that an increase in the overall risk or the level of risk aversion, measured by indices such as VIX, LOIS or EPU, coincides with reductions in FPNRs and lower carry trade returns (risk premium).

Evidence from this study also suggests that FPNRs should be added to the analysis when assessing the factors that destabilize the exchange rate in the short run. The FPNR variable is a plausible candidate for explaining recent strong misalignments in the exchange rate, as identified by Wu (2013). We also recommend considering the asymmetry coefficient (skewness) as an indicator of turbulence in the exchange rate. This variable provides relevant information that is not captured by other commonly used indicators, such as the standard deviation.

From the point of view of monetary policy, the recent rapid development of the carry trade in the case of the Chilean peso is a factor that, in conjunction with other events and economic scenarios, can lead to a situation in which policymakers are faced with a trade-off between inflation control and the incubation of currency risks in the real sector's balance sheets. In this context, hikes in the policy rate intended to curb inflation could spur the carry trade, thereby generating more inflationary pressure. This is a risk that policymakers should take into consideration during periods marked by large interest rate differentials between the Chilean peso and other currencies.

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Annex A1

Figure A1.1
Cross-section of empirical skewness and three-month interest rate differentials
for the period 2002Q1-2012Q1
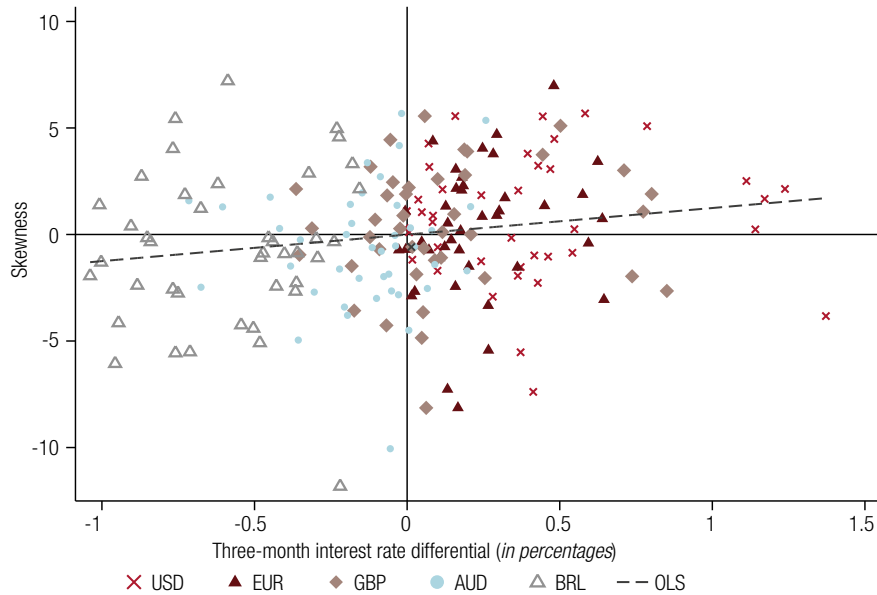
Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.
Note: USD: United States dollars; EUR: Euros; GBP: British pounds; AUD: Australian dollar; BRL: Brazilian reais;
OLS: Ordinary least squares.

Figure A1.2
Cross-section of empirical return and three-month interest rate differentials
for the period 2002Q1-2012Q1
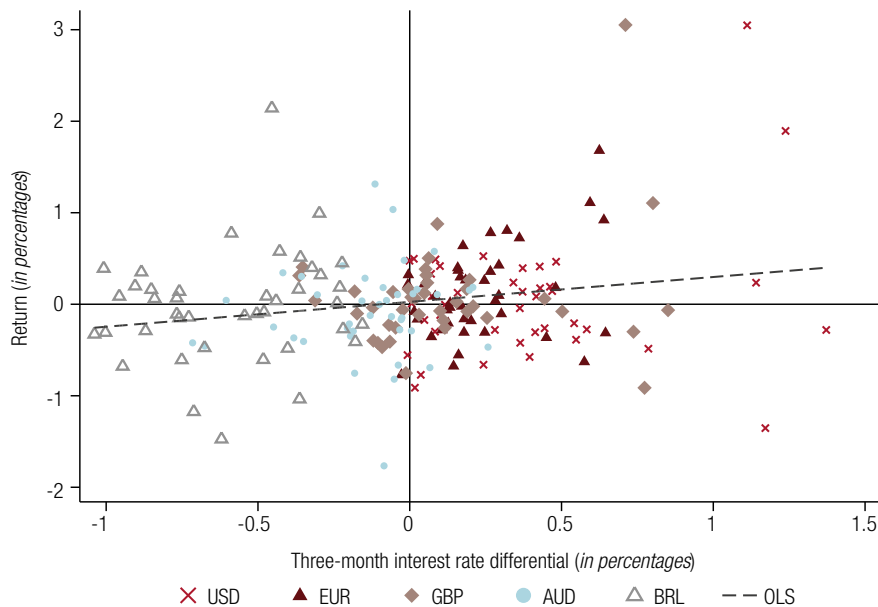
Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.
Note: USD: United States dollars; EUR: Euros; GBP: British pounds; AUD: Australian dollar; BRL: Brazilian reais;
OLS: Ordinary least squares.

Figure A1.3
Three-month interest rate differentials (carry trade) and Chilean peso-United States
dollar exchange-rate volatility
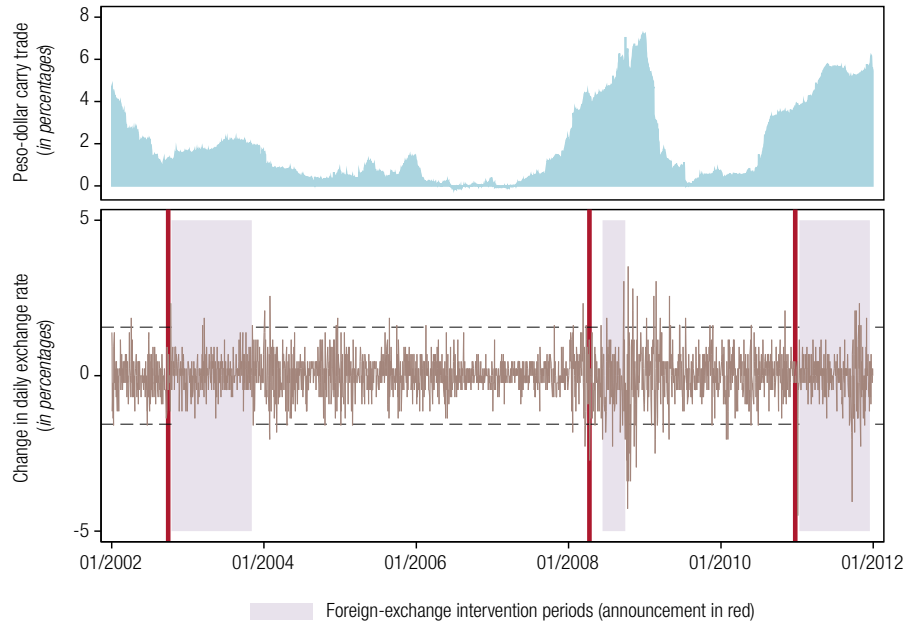
Foreign-exchange intervention periods (announcement in red)

Source: Prepared by the authors, on the basis of information from the Central Bank of Chile.
Note: The red lines represent announcements by the central bank that it would be intervening in the foreign-exchange
            market. The shaded areas mark the periods during which those interventions were implemented. The dashed lines
            in the lower graph correspond to two standard deviation bands (2002-2012 sample).

Figure A1.4
Economic policy uncertainty index, 2002M1-2012M1

--- USD   —— EUR
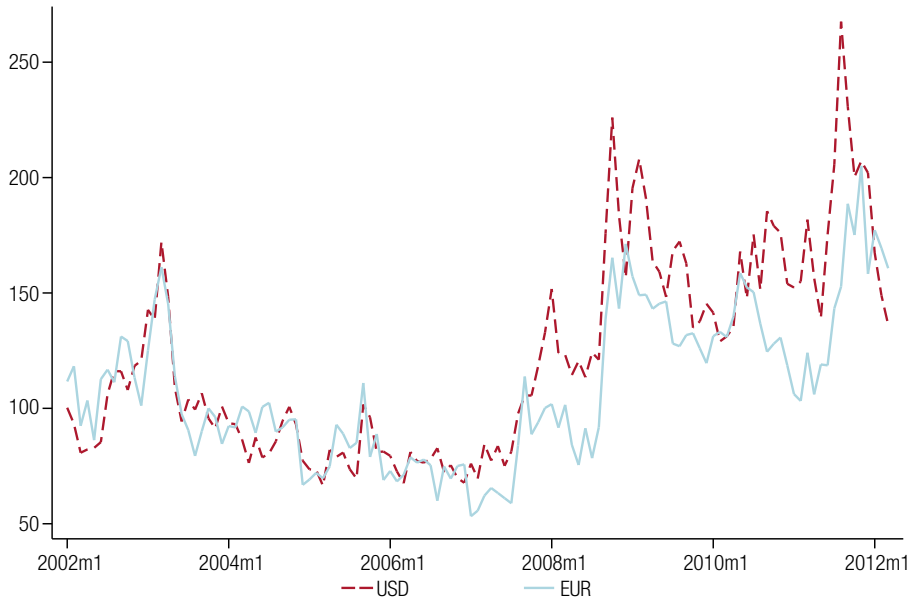
Source: S. Baker, N. Bloom and S. Davis, "Measuring economic policy uncertainty"", *NBER Working Paper Series*, No. 21633,
            Cambridge, Massachusetts, National Bureau of Economic Research, 2015 [online] www.policyuncertainty.com/
            media/BakerBloomDavis.pdf.
Note: USD: United States dollars; EUR: Euros.

Table A1.1
Monthly sensitivity of carry trade positions and returns to changes in the volatility index

	(1) $\Delta FPNR_t$	(2) $\Delta FPNR_{t+1}$	(3) z_t	(4) z_{t+1}
$\Delta VIX_t \times sign(i^*_{t-1} - i_{t-1})$	-12.48	-1.203	-6.46e-06	-0.00106
	(11.82)	(1.396)	(0.000143)	(0.000806)
$\Delta FPNR_{t-1}$	-0.190***	-0.0718***		
	(0.0106)	(0.00567)		
$\Delta log(Copper)_{t-1}$	0.311***	-0.824***		
	(0.0614)	(0.256)		
$BCCH_t$	-64.62	54.48***	-0.00408***	-0.00223
	(45.83)	(9.703)	(0.000141)	(0.00721)
Observations	179	175	183	179
R^2	0.060	0.011	0.004	0.120

Source: Prepared by the authors, on the basis of information from the Central Bank of Chile and data from Bloomberg.
Note: The panel shows country fixed effects and monthly data for the period 2002M3-2012M1 for United States dollars and for the period 2006M8-2012M3 for the euro. The data on forward positions held by non-residents include non-deliverable forward contracts only. Monthly interest rates are shown for the beginning of period t. Clustered standard errors are shown in parentheses: *** $p<0.01$; ** $p<0.05$; * $p<0.1$.

Unemployment dynamics in Uruguay: an analysis using chain reaction theory[1]

Martín Leites and Sylvina Porras

Abstract

This article analyses unemployment dynamics in Uruguay within the framework of chain reaction theory and offers evidence to account for the remarkable drop in unemployment over recent years. It confirms the impact of exogenous variables on the long-run trajectory of unemployment and rules out the notion that its long-run level gravitates around an equilibrium value. It identifies inertia processes in labour supply and demand and in wages, whose mutual interactions mean that shocks have persistent effects on unemployment. There are also complementary spillover effects that influence the magnitude and duration of the effects of shocks. Lastly, the article emphasizes that growth in the capital stock and capital productivity accounts for some of the substantial decline in unemployment in Uruguay since 2003.

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I.   Introduction

The unemployment rate in Uruguay dropped by 12 percentage points between 2002 and 2011, and in recent years has remained at levels that are historically low relative to long-run performance. This suggests a degree of inertia in its dynamics, casts doubt on the existence of a natural rate of unemployment and raises questions about the factors explaining this process. In particular, the matter of how the evolution of long-run unemployment is connected to capital accumulation takes on considerable importance, as this variable became very dynamic in the period, increasing the country's production capacity. Figure 1 shows that in the late 1990s the dynamics of gross fixed capital formation began to move symmetrically relative to the unemployment rate. This suggests the hypothesis that this non-labour market variable could explain some of the drop in the unemployment rate through its positive impact on the demand for labour.

Figure 1
Gross fixed capital formation and unemployment, 1997-2011
(Billions of 2005 pesos and percentages)
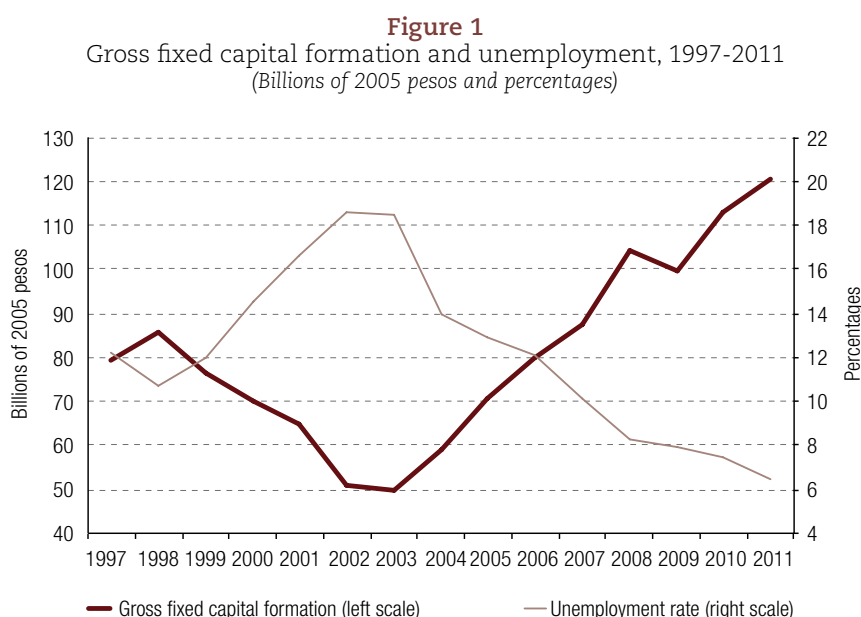
Source: Prepared by the authors, on the basis of data from the National Institute of Statistics (INE) and the Central Bank of Uruguay.

Earlier works studying unemployment dynamics in Uruguay used the most standard approaches in the economic literature: the natural unemployment rate approach and the hysteresis hypothesis. Their findings do not conclusively show the existence of a natural rate of unemployment, and while demonstrating a degree of inertia, they do not provide evidence for the factors determining its evolution. The approaches used in these studies have three essential limitations: (i) they operate with one-equation models; (ii) they do not serve to link short- and long-run behaviour, and (iii) they rule out the introduction of trend variables to explain unemployment.

The purpose of the present study is to explain unemployment dynamics in Uruguay during the period from 1985 to 2011 by applying chain reaction theory, which has some advantages over more extended methods in answering the questions raised. This approach is more general and, instead of assuming that unemployment will converge on an equilibrium level or that the effect of shocks will be persistent (hysteresis), suggests a model for evaluating both hypotheses empirically. Second, it is less restrictive and can be used to identify the effects on the unemployment rate of shocks from variables that are exogenous to the labour market. According to Bande and Karanassou (2009), one of the key advantages of chain reaction theory is that it enables the effects of capital stock changes on the

unemployment rate to be evaluated empirically in both the short and the long run. This is in contrast to the standard approaches, which are emphatic that shifts in the capital trajectory have no long-term effect on the unemployment rate (Layard, Nickell and Jackman, 1991).

Third, it offers a comprehensive perspective connecting the evolution of short-run unemployment to its long-run trend. This makes it possible to identify the adjustment mechanisms operating in the labour market, evaluate the persistence of a temporary shock's effects and analyse how variables exogenous to the labour market, and the interaction of these with supply and demand adjustment processes (spillover effects), can have persistent or permanent effects on long-run unemployment.

To the best of the authors' knowledge, this approach has not been used to study unemployment in developing countries. Certain characteristics of labour markets in the region's countries, such as informality and the level of segmentation, and likewise their remarkable recent dynamism, make it well worth applying (Weller, 2014; World Bank, 2012). Uruguay is no exception, having had historically low levels of unemployment recently by both its own and regional standards, which makes it important to ascertain how unemployment dynamics have been affected by the following: (i) the presence of inertia factors; (ii) variables exogenous to the labour market; (iii) spillover effects, and (iv) the complementarity of adjustment processes. Having evidence on these aspects will contribute to a better understanding of the dynamics of unemployment and make it possible to identify the main determinants of its recent decline. In particular, this study provides evidence for the effects of capital accumulation and labour productivity on the drop in unemployment, variables that Weller (2014) suggests could express the importance of the production context in explaining the dynamism of employment. Furthermore, it indirectly captures the way the production context interacts with institutional aspects, generating mechanisms for transmitting shocks.

The results obtained disprove the notion that the unemployment rate in Uruguay gravitates around an equilibrium level and demonstrate that variables exogenous to the labour market, such as capital accumulation, capital productivity and the working-age population, exert an influence via adjustment processes, generating persistent effects. They also confirm the existence of complementarity in lagged adjustment processes, implying that if these occurred in isolation, over 90% of the initial impact would disappear in the fourth quarter, whereas if they occur simultaneously, this happens only in the tenth quarter.

The article is organized into four sections after this Introduction. First, the contributions of chain reaction theory are reviewed and the results of its application in other countries presented. After that, the methodology applied in this study is described, following which the results are set forth. Lastly, the conclusions of the study are summarized.

II.   Literature review

1.   Chain reaction theory: concept and applications

One of the most influential theories in mainstream macroeconomic studies of unemployment is that there is a natural rate of unemployment or, in the more modern version, a non-accelerating inflation rate of unemployment (NAIRU). Going by evidence on the performance of the labour market in the developed countries over the last three decades of the twentieth century, Layard, Nickell and Jackman (1991) argue for the existence of equilibrium without full employment and identify it as the level of unemployment that does not cause inflation to accelerate. Chain reaction theory arose as an alternative approach to the study of the unemployment rate, dealing with some of the limitations of NAIRU theory: it can be used to model the trajectory of unemployment more comprehensively by combining its short-

and long-term behaviour, evaluate the effects of variables exogenous to the labour market and explain what factors determine the persistence or permanence of temporary shocks.[2]

Chain reaction theory is based on a multi-equation model of the labour market in which changes in the unemployment rate are seen as "chain reactions" to temporary and permanent shocks in that market.

Unemployment reacts via an interacting network of lagged adjustment processes in the major labour market variables: supply, demand and wages. The behaviour of these variables presents a degree of inertia or memory such that the adjustments are not immediate, implying that their current values depend on their past values, with the result that the effects of shocks on the unemployment rate persist. Process inertia is well documented in the literature on chain reaction theory (Karanassou, Sala and Salvador, 2008) and is due to the presence of labour turnover costs (hiring, training and dismissal) and price and wage staggering, and to the existence of internal markets, unemployment duration effects and adjustments in the labour supply (because of emigration, for example). Another important factor is that the processes are interconnected with one another and with exogenous variables, the result being that external shocks create spillover effects, amplifying their short-run impact on the unemployment rate and extending their duration, while preventing the unemployment rate from converging on an invariant level in the long term (Karanassou and Snower, 1997 and 1998).

Empirically, the study of unemployment from a chain reaction theory perspective is approached by modelling a system of equations like the one presented below:[3]

$$n_t = \alpha_1 n_{t-1} + \beta_1 k_t - \gamma_1 w_t \tag{1}$$

$$l_t = \alpha_2 l_{t-1} + \beta_2 z_t + \gamma_2 w_t \tag{2}$$

$$w_t = \beta_3 x_t - \gamma_3 u_t \tag{3}$$

where n, l and w are endogenous variables and represent the demand for labour, the supply of labour and the real wage, respectively.[4] All the variables in the system except the unemployment rate (u) are expressed in logarithms, so that this rate is obtained from the following expression:

$$u_t \cong l_t + n_t \tag{4}$$

The α_i coefficients (where $0 < \alpha_i < 1$) represent the effects of process inertia, while the γ_i parameters reflect the interactions between the endogenous variables (spillover effects)[5] and β_i show the short-run elasticities of the endogenous variables relative to changes in the exogenous variables, such as the capital stock (k), the working-age population (z) and factors that put pressure on prices (x).

This modelling of chain reaction theory deals with one of the main limitations of the NAIRU approach, which assumes that models of the unemployment rate can only include exogenous variables

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[2] Chain reaction theory was developed in Karanassou and Snower (1996). For a review, see Karanassou (1998), Karanassou and Snower (1997, 1998 and 2000) and Henry, Karanassou and Snower (2000), among others.

[3] The theoretical underpinnings for the specification of each equation are set out in Karanassou and Snower (2000). A summary can be found in Leites and Porras (2013).

[4] To simplify, the wage equation excludes the lagged wage effect, following the example of Karanassou, Sala and Snower (2009). Error terms and constants were excluded for the same reason.

[5] In this simple version, wages affect the labour supply and demand, while unemployment also affects wage-setting. Higher levels of unemployment will reduce the relative bargaining power of employed workers by making the threat of dismissal credible, which will tend to hold wages down.

without trend (invariance hypothesis), implying that permanent shocks in exogenous variables lead to compensatory changes in the labour supply and demand curves tending to return unemployment to its original long-run equilibrium rate (Layard, Nickell and Jackman, 1991). A less restrictive kind of modelling (weak invariance hypothesis) includes a stationary linear combination of trend variables that are exogenous to the labour market (Phelps, 1994). Karanassou and Snower (2004) show that these invariance constraints do not apply and that trend variables such as the capital stock, technological change, labour productivity and the working-age population are important in explaining the behaviour of the labour market, and thence unemployment. The requirement is that each endogenous trend variable (l, n and w) should balance out overall with the explanatory variables.

This approach has been applied to the study of the unemployment rate in the United Kingdom, the United States, Denmark, Australia, Portugal, Spain and a group of countries in the European Union (Karanassou and Snower, 1998; Henry, Karanassou and Snower, 2000; Karanassou, Sala and Snower, 2003; Bande, 2002; Bande and Karanassou, 2009; Karanassou, Sala and Salvador, 2008; Karanassou and Sala, 2008 and 2010; González and Sala, 2011).[6] These studies highlight the importance of this approach in comparison with a one-equation model of the unemployment rate, concluding that economic policy decisions should not be based on what the natural rate of unemployment determines, since the unemployment rate is found not to gravitate around this equilibrium level.

Inertia in adjustment processes is confirmed in all cases, with only differences in their speed being identified. The demand for labour adjusts more quickly in countries such as Denmark and Australia ($\alpha_1 = 0.2$ approximately), while at the other extreme is the inertia estimated for the European Union as a whole (0.94), followed by the United States and United Kingdom (about 0.7). It is argued that the finding for Denmark reflects greater flexibility in the country's labour market than in the other countries of the Organization for Economic Cooperation and Development (OECD) in respect of job protection legislation (Karanassou, Sala and Salvador, 2008). Where the labour supply is concerned, the inertia coefficients yielded by the equations are more homogenous and relatively high (α_2 between 0.6 and 0.92),[7] indicating the existence of workforce entry and exit costs. In all cases, the wage equation also includes lags in the dependent variable. Leaving aside the data for Australia and Denmark, the wage inertia coefficients are also relatively homogeneous and range from 0.55 to 0.83. As in the case of labour demand, these are the two countries presenting the fastest adjustment processes (0.32 and 0.24, respectively).

Karanassou and Snower (1998) find that over half the changes in the United Kingdom unemployment rate between 1980 and 1995 were due to the medium-run contribution of lagged adjustments in labour market variables. They also indicate that adjustment processes were complementary, creating more substantial effects on the unemployment rate. Henry, Karanassou and Snower (2000) estimate, for the United Kingdom, that a temporary positive shock in labour demand[8] leads to a gradual reduction in the unemployment rate, so that 90% of the initial effect disappears only after four years. A temporary shock to real wages, meanwhile, leads to a sharp initial increase in the unemployment rate followed by a gradual decline, with 90% of the total adjustment being completed only after 12 years, while it takes 10 years for 90% of the adjustment to a labour supply shock to work through.

The real wage was included as an explanatory variable in the labour demand equations for all the countries. The elasticity of long-run employment relative to wage changes takes the expected sign

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[6] The European Union study included the following countries: Austria, Belgium, Denmark, Finland, France, Germany, Italy, the Netherlands, Spain, Sweden and the United Kingdom.

[7] There are two different estimates for the United Kingdom, one with an inertia coefficient of 0.45 and the other with a coefficient of 0.75.

[8] They assume a shock entailing an initial change of 1 percentage point in the unemployment rate.

in all the countries, fulfilling the law of demand,[9] but the values differ in magnitude. This elasticity takes an absolute value of over 0.6 in one of the estimates for Spain, as well as for the United States and Denmark (where it takes values of between 1 and 2). The other estimates indicate a long-run price elasticity of labour demand of less than 0.5.

Besides the variables mentioned, their lags and their interactions, the equations include other variables exogenous to the labour market as explanatory variables. One of the most-used variables is capital stock, which is included in the labour demand equations. The estimates indicate a positive long-run elasticity ranging from 0.3 to 1.0, implying that there is also a long-run effect on unemployment. In their study of Australia, Karanassou and Sala (2010) conclude that capital accumulation was the factor that most affected the trend of unemployment in the country in the 1990s and early 2000s. On the basis of these findings, the authors question the NAIRU approach and rule out the evolution of unemployment being only a response to temporary shocks or changes in labour market institutions. Karanassou and Sala (2008) find evidence along the same lines for Spain between the 1970s and 2005. They apply chain reaction theory to evaluate the influence of changes in social security contributions, indirect taxes, financial wealth and capital accumulation on the evolution of the unemployment rate in Spain and find capital accumulation to be the factor that does most to explain this.

2. Studies on unemployment in Uruguay

Some earlier studies have addressed the issue of unemployment in Uruguay using the most standard approaches as a framework: the natural rate of unemployment approach and the hysteresis hypothesis. What comes out in these studies is, first, that there is no conclusive empirical evidence for the existence of a natural rate of unemployment on which the Uruguayan economy converges, while in other cases they do not reject the presence of a unit root in the country's unemployment series, or at least do not rule out the possibility that temporary shocks may have persistent effects on unemployment.

Borraz and Tubio (2009) reject the existence of a NAIRU in Uruguay in the period from 1978 to 2009, and while they do find a negative relationship between inflation and unemployment, the method used (expectations-augmented Phillips curve) does not allow them to infer causality.[10] Using a different methodology, though (Kalman filter, univariate modelling), they estimate a "natural" rate of unemployment independent of inflation of 10.64% for 2009, implying that the rate actually recorded between 2009 and 2011 was very far below its "natural" level.

Rodríguez (1998) finds evidence for the hysteresis hypothesis when analysing the evolution of the unemployment rate in Montevideo in the period from 1984 to 1996. Similar results are obtained by Badagián and others (2001), who study the dynamics of the capital's quarterly unemployment rate in the period from the fourth quarter of 1983 to the second quarter of 2001, concluding that unemployment has the behaviour of a long-memory stochastic process (even assuming structural breaks), which would be inconsistent with the existence of a stationary natural rate of unemployment. Another study, also on the Montevideo unemployment rate, in this case during 1968-1997, concludes that the effects of temporary shocks tend to disappear but are highly persistent (Spremolla, 2001).

This evidence implies a degree of inertia in the behaviour of the unemployment rate, which could be explained by long-memory processes in the major labour market variables, such as demand, supply and wage formation. This would imply that there were certain rigidities in the labour market, and that there were no automatic adjustments.

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[9] In the study by Karanassou and Snower (1998) on the United Kingdom, this elasticity presents with a positive sign, but the document does not make any comment on this.

[10] In the authors' judgement, processes driven by independent structural variables are what determine both inflation and unemployment.

The evidence available so far is generally inconclusive. The recent dynamism of the Uruguayan labour market raises the question of what factors might have been determining the low levels of unemployment seen in recent years. Weller (2014) argues that the recent evolution of unemployment in the countries of Latin America and the Caribbean is explained by a favourable production environment (economic growth, rising productivity and the behaviour of other production factors) and by institutional aspects. Approaching this issue via chain reaction theory will provide direct evidence on the importance of factor productivity and capital accumulation in explaining unemployment. It will also provide indirect evidence for the way institutional factors and the production context interact to amplify the effects of shocks.

III. Methodology

A multi-equation dynamic model is estimated in order to explain the evolution of unemployment in Uruguay, taking the chain reaction theory approach as a starting point. Karanassou and Snower (2000) developed the theoretical underpinnings for the modelling of each of the equations in the system.

The labour demand equation is configured to derive this demand from that for final goods and services on the basis of profit maximization for firms operating in non-competitive markets in the presence of adjustment costs. In this framework, demand for labour depends, first, on its own past (lagged dependent variable), which indicates the existence of inertia, owing to the adjustment costs faced by employers in hiring or dismissing workers. For labour demand to be dynamically stable, the coefficient of inertia must be less than 1. Furthermore, the value of this coefficient reflects the speed with which the demand for labour adjusts to external shocks: the closer to 1 the coefficient is, the slower the adjustment will be, and the closer to 0 the coefficient, the quicker the adjustment. The demand for labour also depends on its own price (real wage), the capital stock and capital productivity.

The labour supply equation includes at least one lag of the dependent variable, on the assumption that there are workforce entry and exit costs. This is conducive to inertia in labour supply decisions. Furthermore, changes in supply also depend on other factors: (i) the growth of the working-age population; (ii) real wages, without a predetermined sign, given that supply can react to pay increases positively (substitution effect) or negatively (income effect), and (iii) the unemployment rate, whose effect could be negative insofar as the discouraged worker effect operates (part of the labour force gives up the search for work because of a lack of opportunities) or positive if the added worker effect predominates (more people enter the market in response to an increase in the number of unemployed in their household, in order to make up for the lost income).

Lastly, the wage equation contains at least one lag of the dependent variable, associated with the inertia that is explained by the wage staggering effect. This derives from bargaining between workers and firms, who agree on pay increases in the light of the past trajectory, so that the influence of the lags is immediate. According to Bande (2002), the descriptive model of aggregate wage behaviour is based theoretically on the wage staggering processes described by Taylor (1979). It is assumed that contracts are valid for one period but are determined at two points in time, half at the beginning of the year and the other half in the middle of the period. The result is that the current wage depends on the previous period's wage.[11] Furthermore, wages depend on pay bargaining capacity, which can be modelled as a function of the unemployment level. The size of this variable is believed to capture the insider membership effect on pay bargaining capacity, the idea being that an "insider" group of labour market participants have a privileged position relative to outsiders, explained essentially by the

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[11] Developed in Bande (2002).

cost to firms of staff turnover.[12] Higher unemployment would reduce the bargaining power of insiders and exert downward pressure on wages, while the opposite would happen if unemployment were low. The more traditional theory also includes changes in labour productivity as an explanatory factor in wage movements.[13]

On these foundations, the following system is estimated:

$$n_t = c_n + \alpha_n n_{t-1} + \gamma_1 w_t + \beta_1 k_t + \beta_2 prk_t + \varepsilon_n \tag{5}$$

$$l_t = c_l + \alpha_l l_{t-1} + \gamma_2 w_t + \gamma_3 u_t + \beta_3 z_t + \varepsilon_l \tag{6}$$

$$w_t = c_w + \alpha_w w_{t-1} + \gamma_4 u_t + \beta_4 prn_t + \varepsilon_w \tag{7}$$

Where:

$$u_t = l_t - n_t \tag{8}$$

where n_t is the number of people in work, w_t the average real wage, k_t the capital stock, prk_t apparent capital productivity, l_t the number of people active in the labour market, u_t the unemployment rate, z_t the working-age population and prn_t apparent labour productivity. All the variables other than the unemployment rate are expressed in logarithms, so that the estimated coefficients for each explanatory variable represent the short-run elasticity of the dependent variable relative to changes in these variables. Furthermore, c_i and ε_i (with $i=n$, l and w) represent the constant and the error term of each equation, respectively.

The next step after estimating the system is to contrast the following hypotheses:

- Ha. The presence of an inertia effect. This is considered by analysing the significance of the lagged endogenous variables in each equation. In terms of the model specified, the presence of the inertia effect implies rejection of Ha: $\alpha_n = 0$; $\alpha_l = 0$; $\alpha_w = 0$.

- Hb. The influence of variables exogenous to the labour market. The significance of the following variables is evaluated: the capital stock, capital productivity, the working-age population and labour productivity. Trend variables exogenous to the labour market can be said to affect the unemployment rate if one or more of the following hypotheses hold true: $\beta_1 \neq 0$; $\beta_2 \neq 0$; $\beta_3 \neq 0$; $\beta_4 \neq 0$.

- Hc. Spillover effects. Within the framework of the model specified, there are found to be spillover effects if one or more of the following hypotheses hold true: $\gamma_1 \neq 0$; $\gamma_2 \neq 0$; $\gamma_3 \neq 0$; $\gamma_4 \neq 0$. This would imply that there was interaction between the equations, so that a shock found to affect one of the endogenous variables will create spillover effects on the others, which could result in the unemployment adjustment process being even more prolonged.

The above hypotheses were contrasted by evaluating the statistical significance of the coefficients estimated in each equation. According to chain reaction theory, non-rejection of hypotheses Ha and Hc implies that temporary shocks will have a persistent effect on the unemployment rate. Simulation techniques based on impulse-response functions were used to evaluate this persistence.

- Hd. Hypothesis of complementarity in lagged adjustment processes. Another of the aspects flagged up in the literature on the subject is that the effects on unemployment can be amplified

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[12] Montuenga and Ramos (2005) propose different justifications for bringing in the unemployment rate as a determinant of wages.

[13] Efficiency wage models suggest that labour productivity could be endogenous to the system, an aspect that has not hitherto been considered in chain reaction theory.

and/or prolonged when there is complementarity in adjustment processes. As a way of contrasting this hypothesis in the case of Uruguay, we follow Henry, Karanassou and Snower (2000) and use simulation techniques to compare the effects of a shock on unemployment in the following two simulations: one in which all the effects are considered to operate simultaneously, and one that consists in adding up the individual influences of each equation as though they operated separately.

Lastly, a secondary objective is to investigate the determinants of the recent drop in unemployment in Uruguay, exploring how important capital accumulation has been for the unemployment dynamics of the past few years. For this, following Karanassou and Snower (1998), a simulation exercise is carried out in order to identify the factors contributing most to the drop in unemployment between 2003 and 2011.

We estimated the equations using autoregressive modelling with distributed lags of order p and q, analysing the cointegration of the variables included in the modelling with the methodology proposed in the studies by Pesaran and Shin (1995), Pesaran, Shin and Smith (1996 and 2001), and Pesaran (1997), with p and q corresponding to the order of lags of the dependent variable and the independent variable, respectively.

This methodology is the one used in the earlier literature applying the chain reaction theory approach, as it presents certain advantages over the cointegration techniques normally employed to estimate long-term relationships and short-term adjustment mechanisms. In the first place, it is useful for evaluating the significance of the coefficients of the lags of the endogenous variables, which has a clear economic interpretation in the chain reaction theory framework. Second, this approach does not require a priori knowledge of the order of integration of the variables to analyse long-term relationships, which means that the problems involved in identifying unit roots do not have to be dealt with. Also avoided as a result are the problems associated with the application of traditional cointegration methods, which require long series.

Since the lagged dependent variable is included in the modelling of each equation, estimating the system by ordinary least squares (OLS) could present problems of endogeneity and correlation with residuals. To mitigate these problems, instrumental variables are employed and the model is estimated using the three-stage least squares (3SLS) procedure.

Quarterly series were used for the period between the first quarter of 1985 and the fourth quarter of 2011 (see table A1.1 of the annex). The labour market information used was arrived at by processing microdata from the continuous household surveys of the National Institute of Statistics (INE) and is representative of the urban population of Uruguay (settlements of 5,000 inhabitants and over).[14] The series for the working-age urban population was constructed from the urban population projections of INE and the Latin American and Caribbean Demographic Centre (CELADE)-Population Division of the Economic Commission for Latin America and the Caribbean (ECLAC). The real wage series comes from INE wage statistics. The capital stock series was constructed by turning an annual series available in Román and Willebald (2012) into a quarterly one using a constant depreciation rate and quarterly investment information from the national accounts of the Central Bank of Uruguay.

Two apparent productivity series were constructed to gain an idea of the evolution of capital and labour productivity. The first was obtained from the ratio between gross domestic product (GDP) as reported by the central bank and the capital stock. This is a proxy for apparent capital productivity, since the ratio incorporates capital stock rather than capital use, which would be the correct choice. Apparent labour productivity was constructed from the ratio between GDP and total hours worked, a figure obtained by processing the continuous household surveys.

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[14] This information is used because the survey was extended to smaller settlements and rural areas only in 2006.

IV. Results

1. Estimating the system of equations

The econometric methodology applied first requires the F-test to be contrasted with the three equations making up the multi-equation dynamic model: labour demand, labour supply and wages. This test evaluates the null hypothesis (H_0); namely, that there is no cointegration between the variables. In all three cases, the F-statistic was above the critical value in the tables given by Pesaran, Shin and Smith (2001), so that it was possible to reject H_0 and conclude that there was a long-term relationship between the variables included in each model (see table A1.2 of the annex).

With the existence of a long-term relationship confirmed, the multi-equation dynamic model was estimated. The results discussed below are those obtained by estimating the system of equations using the 3SLS method (see table 1).[15]

Table 1
Estimation of labour market equations by the three-stage least squares (3SLS) method[a]

Labour demand n_t			Labour supply l_t			Wages w_t		
Variable	Coefficient	p-value	Variable	Coefficient	p-value	Variable	Coefficient	p-value
n_{t-1}	0.876	0.00	l_{t-1}	0.774	0.00	w_{t-1}	0.786	0.00
w_{t-4}	-0.055	0.00	u_{t-1}	-0.086	0.02	prn_{t-1}	0.095	0.00
k_{t-1}	0.056	0.01	Δw_t	0.093	0.01	Δprn_{t-1}	0.053	0.03
prk_{t-1}	0.081	0.00	z_{t-1}	0.336	0.00	u_{t-1}	-0.441	0.00
Δprk_t	0.092	0.00						
d1	-0.042	0.00	d3	-0.036	0.00	d1	-0.073	0.00
d2	-0.035	0.01	d4	0.016	0.06	d6	-0.031	0.00
d3	-0.061	0.00	d2	-0.037	0.00	d7	-0.046	0.00
C	2.163	0.00	d5	0.019	0.02	d8	0.041	0.01
			C	-1.714	0.00	d9	0.043	0.00
						C	0.594	0.00
Number of observations	103			103			103	
R²	0.98			0.98			0.97	
SSR[b]	0.02			0.02			0.02	
Joint test for normality (p-value)				0.11				
Joint test for autocorrelation (p-value)								
			1 lag	0.16				
			2 lags	0.50				
			3 lags	0.64				
			4 lags	0.57				

Source: Prepared by the authors.

[a] Instrumental variables: n_{t-1}, l_{t-1}, w_{t-1}, w_{t-4}, w_{t-5}, k_{t-1}, k_{t-2}, prk_{t-1}, prk_{t-2}, Δprk_t, Δprk_{t-1}, u_{t-1}, u_{t-2}, u_{t-3}, u_{t-4}, Δw_t, Δw_{t-1}, Δw_{t-2}, z_{t-1}, z_{t-2}, z_{t-3}, prn_{t-1}, prn_{t-2}, Δprn_{t-1}, Δprn_{t-2}, d1, d2, d3, d4, d5, d6, d7, d8, d9 and c.

[b] Sum of squared errors.

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[15] The results obtained with the 3SLS method do not differ substantially from those produced when the system is estimated by OLS. See table A1.5 of the annex in Leites and Porras (2013).

The estimated coefficients of the variables included in the three equations of the system proved to be significant, with the expected signs; in addition, their residuals are well behaved (normal, without autocorrelation or heteroskedasticity), both in the estimation of the system and in the estimation of each equation separately (see table A1.3 of the annex).

The results indicate that the demand for labour depends, in the first place, on its own past (n_{t-1}), which confirms the existence of inertia due to the adjustment costs to employers of hiring or dismissing workers, such as the cost associated with training up new employees. The estimation yields a labour demand inertia coefficient of 0.876, a fairly high level by the standards of the estimates for other countries that were mentioned earlier, indicating that labour demand shocks will have effects on the unemployment rate that will not quickly disappear.

In addition, as was to be expected, the demand for labour relates negatively to its price (the real wage), but this relationship does not link the two elements simultaneously, as movements in the real wage impact labour demand only after a four-quarter lag. According to the coefficient estimated, an increase of 1% in real wages reduces long-run labour demand by 0.45% [= -0.055/(1-0.876)]. It should be noted that the real wage variable refers to the cost of employees, so that presumably the elasticity estimated would be greater if the universe of demand were restricted to workers of this type.

As in other countries, the capital stock in Uruguay explains labour demand, in this case with one lag. A rise of 1% in the capital stock generates a total rise of 0.45% [= 0.056/(1-0.876)] in labour demand. This elasticity is close to the values estimated for a number of the countries, as already discussed. Not only did the capital stock prove significant in explaining shifts in labour demand, but so did capital productivity, in level and variation. The long-run elasticity of labour demand relative to the level of capital productivity was estimated at 0.65, while the long-run elasticity of that demand relative to the variation in capital productivity was estimated at 0.74.

The estimated inertia coefficient in the labour supply equation (0.774) is rather lower than the one for labour demand, but within the range of values estimated for the other countries. As with the countries where the unemployment rate was included as an explanatory factor for supply, the sign of the elasticity indicates that the discouraged worker effect also operates in Uruguay. What this means is that an increase of 1 percentage point in the unemployment rate translates into a drop of 0.38% in the labour supply [= -0.086/(1-0.774)],[16] indicating a widespread belief among people seeking work in this context that it is going to be difficult to find, so that a proportion of them abandon the search and drop out of the market. At the same time, the labour supply reacts positively (0.41%) [= 0.093/(1-0.774)] to a 1% rise in the real wage, indicating that the substitution effect prevails over the income effect, i.e., that when wages increase, so does the number of people preferring to substitute work for leisure. Lastly, as would be expected, the working-age population is the exogenous variable that does most to explain the evolution of the labour supply. Thus, an increase of 1% in this variable leads to a rise of 1.49% in the supply [= 0.336/(1-0.774)], indicating an elasticity that is around the average of the values estimated for developed countries.

The significance of the lagged dependent variable in the wage equation shows that real wages also evince a degree of inertia (0.786) associated with the wage staggering effect. The size of this coefficient is close to the values estimated for most developed countries, which indicates in this case too that wage shocks will have effects on the unemployment rate that do not disappear quickly. For its part, the coefficient for the unemployment rate that was included in this equation could be capturing a degree of flexibility in the response of real wages to labour market conditions, in the sense that employed workers lose wage bargaining power when unemployment is rising, which places downward pressure on real wages. Thus, a rise of 1 percentage point in the unemployment rate leads to a long-

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[16] The dependent variable is logarithmic, whereas the unemployment variable is a rate, so that the estimated coefficient multiplied by 100 is interpreted as the semi-elasticity of the labour supply relative to the unemployment rate.

term reduction of 2.06% in real wages. Although having the same sign, this elasticity is slightly above the values (ranging from 0.5% to 1.5%) estimated for the other countries whose supply equations included this variable as an explanatory factor for real wages. Wages also respond to changes in the levels and variation of labour productivity per hour worked. The long-run elasticity of wages was estimated at 0.44 (total effect) relative to labour productivity levels and 0.25 relative to variation. These values are below those found for the United States in 2011 using a comparable specification (1.0 and 3.29, respectively).

Dummy variables also had to be included in all three equations, usually because of outlying one-off values in the series or events in the Uruguayan economy that produced some break in the series, such as the 2002 crisis (see table A1.1 of the annex).

The system estimated yields a satisfactory prediction of the evolution of the unemployment rate in Uruguay over the last 27 years. This can be seen in figure 2, which presents the actual evolution of the unemployment rate $(l_t - n_t)$ and the evolution estimated from the system of equations (with and without dynamic interaction).[17]

Figure 2
Evolution of the actual unemployment rate and the rate estimated from the system
of equations, 1985-2011
(Percentages)

| | Actual | | Estimated (with dynamic interaction) | | Estimated (without dynamic interaction) |
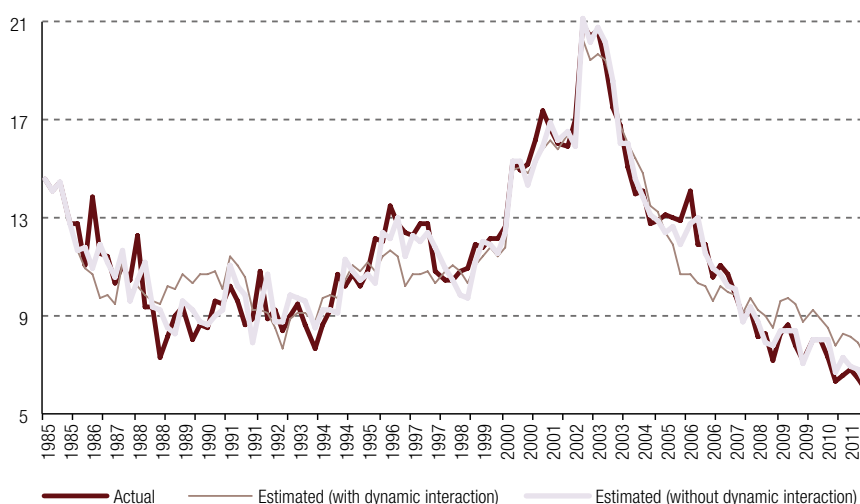
Source: Prepared by the authors.

2. Hypothesis testing and implications

The results of the estimates provide a response to the first three questions raised. Table 2 presents the hypotheses and the statistical tests and summarizes their results. In the first place, all three equations confirm the significance of the lagged dependent variables $(\alpha_i \neq 0)$, which is evidence for inertia in labour market adjustment processes (Ha). In the second place, it is found that the capital stock and capital productivity have a significant effect on labour demand, that the working-age population affects the labour supply and that labour productivity affects wages, confirming the influence of exogenous variables on the unemployment rate $(\beta_i \neq 0, Hb)$. Lastly, wages are important in explaining labour

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[17] The inclusion of dynamic interaction means that the unemployment rate estimated by the system is then incorporated into successive estimates of labour supply and wages. The projection that does not incorporate dynamic interaction takes the actual unemployment rate as an explanatory variable for supply and wages.

supply and demand, and unemployment is important in explaining both wages and the labour supply, thus confirming the significance of interactions between the equations and the presence of spillover effects ($\gamma_i \neq 0$, Hc).

<div align="center">

Table 2
Hypothesis testing[a]

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Hypothesis		Equations[b]			Conclusion
		Demand	Supply	Wages	
Ha: presence of inertia effect H0: coefficient = 0 H1: coefficient ≠ 0	Estimated coefficient	$\alpha_n = 0.876$	$\alpha_l = 0.774$	$\alpha_w = 0.786$	Presence of inertia not rejected
	t-test (p-value)	0.00	0.00	0.00	
Hb: influence of exogenous variables H0: coefficient = 0 H1: coefficient ≠ 0	Estimated coefficient	$\beta_1 = 0.056$	$\beta_4 = 0.336$		Influence of exogenous variables not rejected
	t-test (p-value)	0.01	0.00		
	Estimated coefficient	$\beta_2 = 0.081$			
	t-test (p-value)	0.00			
	Estimated coefficient	$\beta_3 = 0.092$			
	t-test (p-value)	0.00			
Hc: presence of spillover effects H0: coefficient = 0 H1: coefficient ≠ 0	Estimated coefficient	$\gamma_1 = -0.055$	$\gamma_2 = -0.086$	$\gamma_4 = -0.441$	Presence of spillover effects not rejected
	t-test (p-value)	0.00	0.02	0.00	
	Estimated coefficient		$\gamma_3 = 0.093$		
	t-test (p-value)		0.01		

Source: Prepared by the authors.

[a] Contrasts of hypotheses based on the results of the estimations of the system equations presented in table 1.

[b] The equations are as follows:
Demand: $n_t = \alpha_n n_{t-1} + \beta_1 k_{t-1} + \beta_2 prk_{t-1} + \beta_3 \Delta prk_t + \gamma_1 w_{t-4}$
Supply: $l_t = \alpha_l l_{t-1} + \beta_4 z_{t-1} + \gamma_2 u_{t-1} + \gamma_3 \Delta w_t$
Wages: $w_t = \alpha_w w_{t-1} + \beta_5 prn_{t-1} + \beta_6 \Delta prn_t + \gamma_4 u_{t-1}$
Where:
n: Number of people in work; w: Average real wage; k: Capital stock; prk: Apparent productivity of capital; l: Number of people active in the labour market; u: Unemployment rate; z: Working-age population; prn: Apparent productivity of labour.

The confirmation of these three hypotheses suggests that a temporary shock to any of the labour market variables can generate a persistent effect on the level of unemployment. To evaluate the magnitude of these effects, exercises were first carried out to simulate demand, supply and wage shocks. This made it possible to measure the persistence in time of their effects on the unemployment rate. Second, shocks in variables exogenous to the labour market were simulated. Lastly, this was followed by an evaluation of lagged adjustment processes to ascertain whether they were complementary and amplified their effects (hypothesis Hd).

(a)   Demand, supply and wage shocks

Two types of shocks to the labour demand, wage-setting and labour supply equations were simulated: a one-off shock and an autoregressive shock of type AR(1) with $\rho = 0.5$. The effects of the shocks were normalized so that the initial impact on the unemployment rate would be 1 percentage point.

As can be seen in figures 3A and 3B, shocks generate different dynamics for unemployment depending on whether they originate from demand, supply or wages. One-off demand and supply

shocks lead to an increase in unemployment,[18] with effects that persist for 11 and 8 quarters, respectively. The effects of the shock diminish by over 90% by the end of the ninth quarter in the first case and the sixth quarter in the second. Wage shocks have a less substantial immediate effect, but their effect on unemployment becomes more significant after a year, as higher wages negatively affect the demand for labour with a lag of four quarters, and the effects persist for over five years.

Figure 3
Trajectory of unemployment in response to labour demand, labour supply and wage shocks
(Percentage points)

A. One-off shock

B. Shock of type AR(1) with ρ = 0.5

Persistence of effect in quarters

Persistence of effect in quarters

———— Demand shock   ———— Supply shock   ———— Wage shock
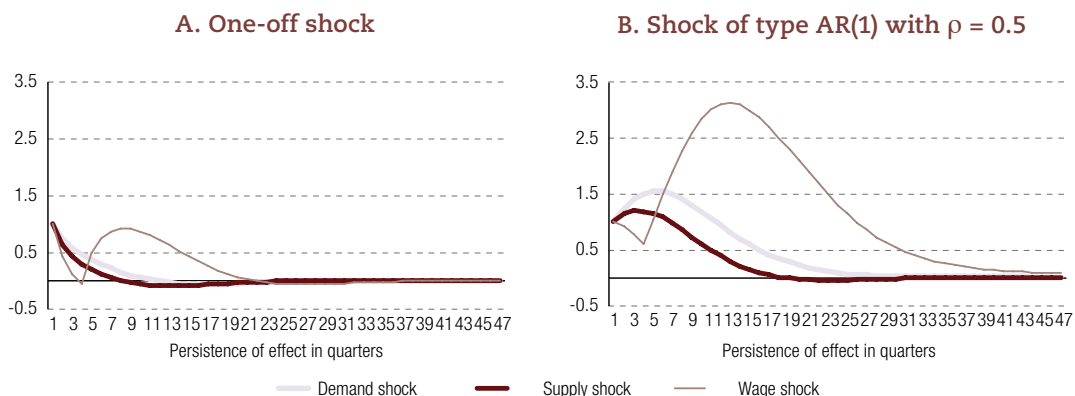
Source: Prepared by the authors.

In response to AR(1) shocks, unemployment overshoots and is highly persistent in the long run (the effects take roughly 5 to 13 years to disappear). When it comes to wages, unemployment overshoots more, rising practically without check until well into the third year, followed by effects that disappear only very slowly. Once again, it was in the study on the United Kingdom that the effects of wage shocks on unemployment were found to be most persistent relative to those of demand or supply shocks (Henry, Karanassou and Snower, 2000).

(b) Shocks to exogenous variables

This exercise illustrates how the interaction of exogenous variables along with delayed adjustments generates persistent effects on the long-run unemployment rate.

Three types of shocks were simulated for each exogenous variable: a one-off shock (without memory) and two AR(1) shocks, with ρ = 0.5 and ρ = 0.9. In all cases, the effects were normalized so that the initial impact on unemployment was a change of 1 percentage point in the rate. The results are presented in figures 4A to 4D.

In the case of shocks to the capital stock, capital productivity and the working-age population, unemployment is found to follow a similar trajectory. With a one-off shock, unemployment reacts and then gradually returns to its starting level; when the shock has memory, however, unemployment initially overshoots, with the initial effect sometimes being more than doubled before it very slowly dissipates. When a shock derives from the capital stock, the effects can be seen to persist for up 10 quarters in the case of a one-off shock and up to 25 in the case of one with lags.

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[18] What is assumed in this case is a negative labour demand shock leading to increased unemployment.

Figure 4
Effects on the unemployment rate of shocks to exogenous variables
(Percentage points)

A. Capital stock

Persistence of effect in quarters

B. Capital productivity

Persistence of effect in quarters

C. Working-age population

Persistence of effect in quarters

D. Labour productivity

Persistence of effect in quarters

—— One-off shock   —— AR(1) shock with $\rho = 0.5$ —— AR(1) shock with $\rho = 0.9$
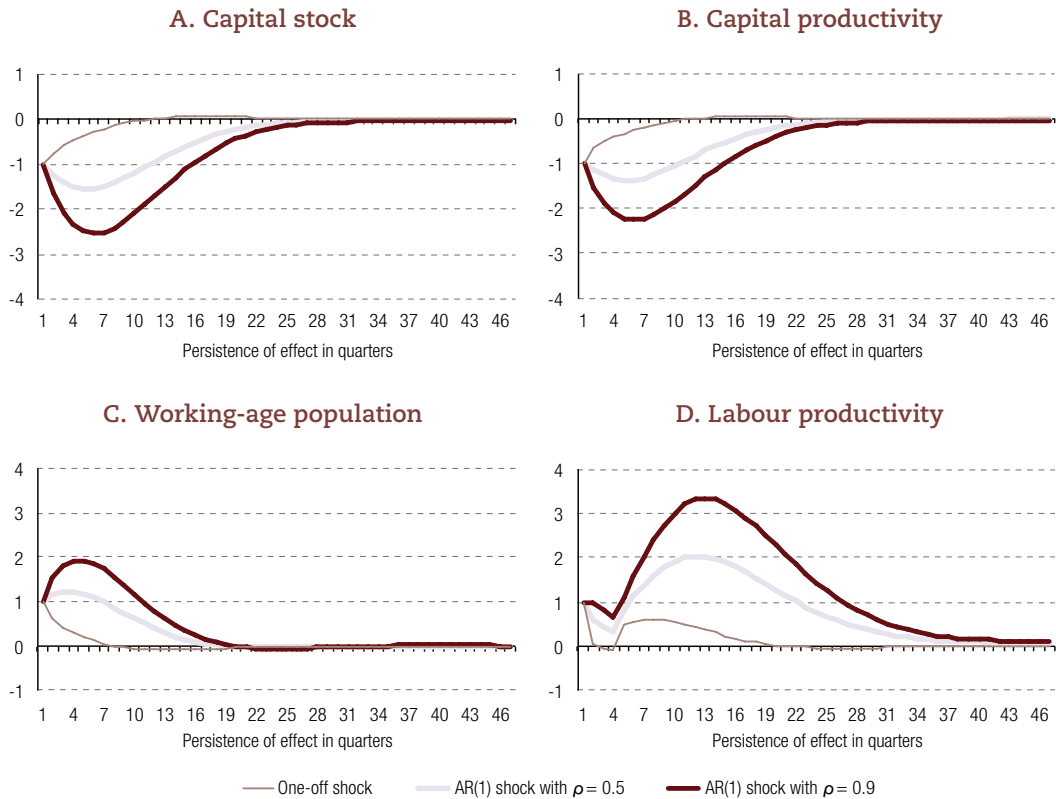
Source: Prepared by the authors.

The effects of labour productivity shocks on unemployment do not present such a gradual trajectory as in the previous cases, as there is first an immediate drop in unemployment and then an overshoot due to the demand effects of the lagged wage. The effect on unemployment is much greater in subsequent periods than initially, peaking at 3.34 percentage points, and is also more persistent (it only disappears after a period that may range from 5 to 17 years, depending on the type of shock).

(c)   Process complementarity

To gauge the magnitude of the complementarities in the adjustment processes empirically, we follow Henry, Karanassou and Snower (2000). A labour demand shock is assumed and the results of the following two simulations are compared: one assuming that all the effects operate simultaneously, and one arrived at by adding up the individual contributions of each equation.

To derive the individual contribution of the demand equation to the unemployment adjustment process, all the endogenous variables of the system are first fixed at their current value, with the exception of the employment lag in the demand equation. In this way, the shock simulation will only encompass the effect on the unemployment rate of inertia in the demand for labour itself. Next, the individual contribution of the supply equation in the face of this demand shock is obtained, this time by fixing all the endogenous variables at their current value except the one for the supply lag. The same procedure is followed to obtain the individual contribution of the wage equation.

In the next step, the individual responses are added together and normalization is carried out so that the immediate effect of the shock on unemployment is 1 percentage point. This yields a series for the effect of a demand shock on unemployment when the adjustment processes act in isolation.

Figure 5 shows that a demand shock has an amplified effect on unemployment when adjustment processes are complementary, and that this takes longer to disappear, providing evidence for hypothesis Hd. In the case of a one-off shock, the effects disappear only after 11 quarters, whereas if adjustment processes did not complement one another their effects would disappear in less than half the time. If the shock additionally follows an AR(1) process with $\rho = 0.5$, complementarity means not only that its effects are longer-lasting, but that they are amplified, even resulting in an overshoot.

Figure 5
Complementarity of lagged adjustment processes: individual and joint effects of a labour demand shock on the unemployment rate[a]
(Percentage points)

A. One-off shock

B. AR(1) shock with $\rho = 0.5$

Persistence of effect in quarters

Persistence of effect in quarters

——— Joint effect   ——— Sum of individual effects
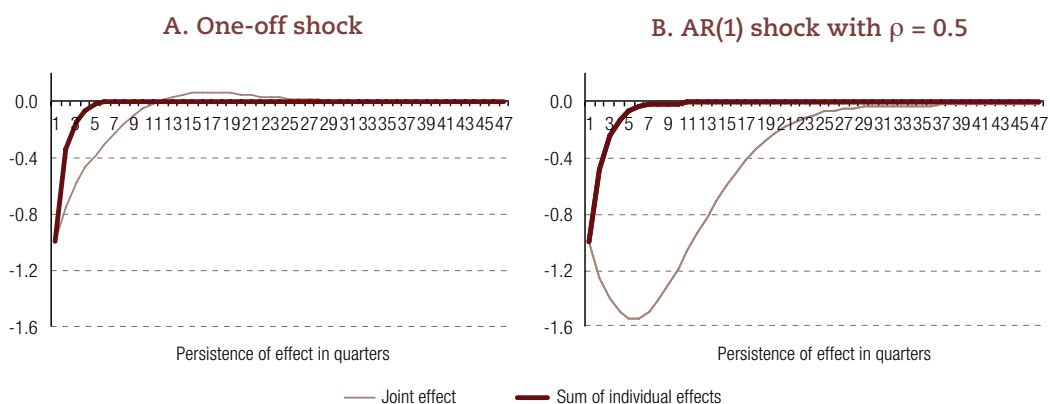
Source: Prepared by the authors.

[a] The effects are calculated by means of two simulations: one assuming that they all operate simultaneously and including their interactions (joint effect), and one arrived at by adding up the individual contributions of the demand, supply and wage equations to the unemployment adjustment process.

3. Determinants of the recent evolution of unemployment

Lastly, an objective of this article was to attempt to ascertain which factors accounted for the drop in the Uruguayan unemployment rate from the high levels of the 2002 crisis to the historically low figures of recent years. To this end, the analysis concentrated on the period from the first quarter of 2003 to the fourth quarter of 2011. It first distinguished the changes in the unemployment rate over the period, analysing the contribution of inertia in adjustment processes. Secondly, it considered what contribution had been made by the complementarity of adjustment processes. Lastly, it explored the influence of exogenous variables on the trajectory of unemployment during the period.

Following Karanassou and Snower (1998), we estimated the contribution of inertia processes to the drop in the unemployment rate ($\Delta u\hat{}_{2011-2003}$). For this, we first calculated the total change in the unemployment rate between 2003 and 2011, using the full system of equations ($\Delta u_{2011-2003}$). The result was that, with all the adjustments in place, the unemployment rate dropped by 10.69 percentage points in the period. We then calculated this same change, but under the assumption that the lagged adjustments of the endogenous variables did not operate in the system ($\Delta u^n_{2011-2003}$). These variables were assumed to stand at their current level, meaning that the adjustment process was completed in each period. This procedure yielded a 5.62 percentage point drop in the rate and made it possible to measure what the change in unemployment would have been if the changes in the exogenous variables alone had operated. The difference between the two movements yields the

contribution of the inertia and the interactions, accounting in this case for almost half the drop in the unemployment rate ($\Delta u^{\wedge}_{2011-2003} = \Delta u_{2011-2003} - \Delta u^n_{2011-2003} = 5.07$ percentage points) (see table 3), with the rest being accounted for by the effects of the exogenous variables.

Table 3
Contribution of lagged adjustments to the change in the unemployment rate
between 2003 and 2011
(Percentage points)

			Change in the unemployment rate
Total	$\Delta u_{2011-2003}$	=	-10.69
Without lagged adjustments in endogenous variables	$\Delta u^n_{2011-2003}$	=	-5.62
Joint contribution of lagged adjustments	$\Delta u^{\wedge}_{2011-2003}$	=	-5.07
With demand inertia only	Δu (EA)$_{2011-2003}$	=	-6.25
With wage inertia only	Δu (WF)$_{2011-2003}$	=	-5.61
With supply inertia only	Δu (LF)$_{2011-2003}$	=	-4.93
Individual inertia contributions			
$\Delta u^{\wedge}(EA)_{2011-2003}$ = Δu (EA)$_{2011-2003}$ - $\Delta u^n_{2011-2003}$		=	-0.63
$\Delta u^{\wedge}(WF)_{2011-2003}$ = Δu (WF)$_{2011-2003}$ - $\Delta u^n_{2011-2003}$		=	0.01
$\Delta u^{\wedge}(LF)_{2011-2003}$ = Δu (LF)$_{2011-2003}$ - $\Delta u^n_{2011-2003}$		=	0.69
Joint contribution of individual effects			
$\Delta u^{\wedge}(EA)_{2011-2003} + \Delta u^{\wedge}(WF)_{2011-2003} + \Delta u^{\wedge}(LF)_{2011-2003}$		=	0.07

Source: Prepared by the authors.
Note: Δu: Change in projected unemployment; Δu^n: Change in unemployment without inertia; $\Delta u^{\wedge}$: Change in unemployment due to inertia; EA: Demand inertia; LF: Supply inertia; WF: Wage inertia.

To measure the importance of the complementarity of lagged adjustment effects, a simulation was made of the change in the unemployment rate between those years under the assumption that this complementarity does not exist. To estimate the change without adjustment complementarity, the contributions of the individual inertia of each variable are added together. In the case of labour demand ($\Delta u^{\wedge}(EA)_{2011-2003}$), for example, the individual contribution of inertia is given by the difference between the change in the unemployment rate assuming demand inertia only ($\Delta u(EA)_{2011-2003}$) and the change in the unemployment rate without any inertia process ($\Delta u^n_{2011-2003}$). Similarly, the contribution of inertia in the labour supply ($\Delta u^{\wedge}(LF)_{2011-2003}$) and in wages ($\Delta u^{\wedge}(WF)_{2011-2003}$) is estimated. As can be seen in table 3, the unemployment rate would have risen by 0.07 percentage points instead of dropping by 5.07 percentage points as described above. It can be observed that the drop in the unemployment rate attributable to demand inertia between 2003 and 2011 was 0.63 percentage points, while that attributable to supply inertia was 0.69, but the contribution of wages was almost nil.

Lastly, what would have happened to the unemployment rate if the exogenous variables had not registered changes between 2003 and 2011? The capital stock and capital productivity, the working-age population and labour productivity all rose between those years. The exercise is to simulate the trajectory of unemployment by altering the value of the exogenous variables.

As a reference, we considered the series for the unemployment rate produced by estimating the full system of equations. After this, the contribution of the capital stock to the evolution of the unemployment rate was calculated by simulating the trajectory of unemployment on the assumption that the capital stock in each quarter of the period from 2004 to 2011 was at the same level as in the equivalent quarter of 2003, with the estimation taking the actual values of the other exogenous variables. The contribution of this variable was obtained as the difference between the two series. The same procedure was applied for all the exogenous variables.

Figure 6 presents the results. When the differences take positive values, it means that the unemployment rate would have been higher if the exogenous variable had kept its 2003 values and not evolved as it did. According to this exercise, the increases in capital productivity and capital accumulation made a substantial contribution to the recent trajectory of unemployment. This means that if the capital stock and capital productivity had not moved upward but had held steady at their 2003 levels, the unemployment rate would have fallen by less than it did, ending the period 3.7 and 7.5 percentage points above the level actually attained, respectively.[19]

Figure 6
Contribution of exogenous variables to the drop in the unemployment rate, 2003-2011
(Percentage points)
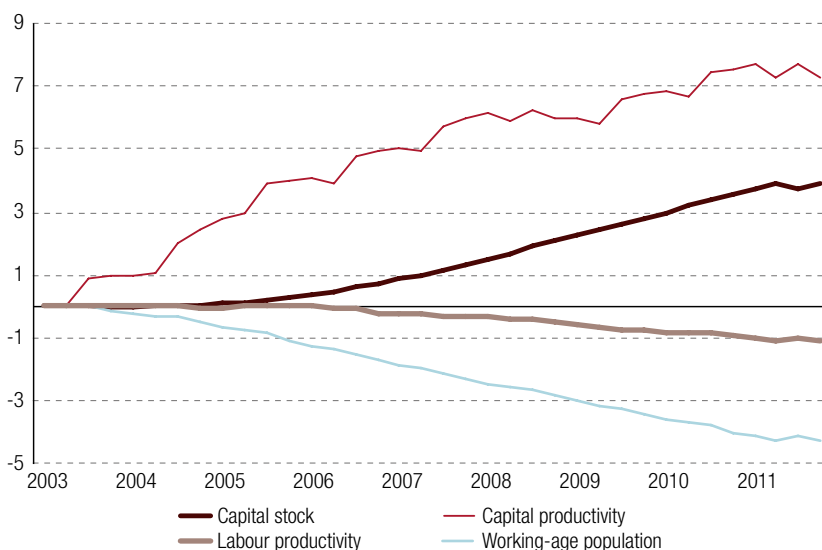
Source: Prepared by the authors.

The other variables had the opposite effect. The increase in the working-age population led to a rise in the labour supply, placing upward pressure on unemployment. If this variable had kept its 2003 level, the unemployment rate in 2011 would have been 4 percentage points lower. Labour productivity was another variable that rose in the period and placed direct upward pressure on wages even as it indirectly influenced demand and supply changes, so that it also exerted upward pressure on unemployment, albeit to a much lesser extent than the rise in the working-age population.

In summary, it can be said that the developments illustrated by figure 6 are consistent with expectations. It shows that the growth in capital stock and capital productivity in Uruguay, combined with the complementarity of labour market adjustments, can explain much of the substantial drop in the unemployment rate between 2003 and 2011.

V. Conclusions

Chain reaction theory was applied to study the dynamics of unemployment in Uruguay in the period from 1985 to 2011, and the evidence supplied helps to explain the determinants of the unemployment level and the decline in the unemployment rate from 2003 onward.

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[19] The level actually attained is not the one actually observed, but the one arrived at by estimating the system as a whole.

First, the major labour market variables, namely demand, supply and wage formation, were found to present inertia and, being interconnected, spillover effects. The result is that one-off shocks to any of the labour market variables have effects on unemployment that take time to dissipate. For example, if the shock is from demand or supply, there are effects on unemployment that persist for 11 quarters and 8 quarters, respectively, while if the shock is from wages, the effects linger for up to 5 years.

Second, adjustment processes operate in a complementary fashion, creating effects whereby the magnitude and persistence of shocks expand. If there were no such complementarity, a one-off shock that increased the demand for labour by causing a 1 percentage point decline in the unemployment rate would dissipate in the fifth quarter, whereas in fact, as indicated, the effect lingers for 11 quarters.

Furthermore, one of the key contributions of chain reaction theory is that it can be used to detect the influence on the unemployment rate of changes in variables exogenous to the labour market. In particular, it was found that changes in the capital stock created persistent effects on unemployment, with these dissipating only after 3 to 7 years, depending on the type of shock.

With regard to the recent performance of unemployment in Uruguay, the reduction in unemployment levels was found to be accounted for essentially by two factors. First, among the exogenous variables, the most important effects were from the rise in capital productivity and the capital stock. If these variables had remained at their 2003 levels, the unemployment rate would have fallen by less, ending the period 3.7 and 7.5 percentage points, respectively, above the level actually attained. Second, a large proportion of the decline is explained by the spillover effects arising from the complementarity of the adjustment processes of Uruguayan labour market variables.

Lastly, the findings of this study have some important implications for both unemployment dynamics and the design of future policies. First, the inertia and complementarity found in the adjustment processes of labour market variables, together with the existence of spillover effects with persistent impacts, suggest that the unemployment rate does not converge on an invariant level. On the contrary, its long-term evolution appears to be determined by the unending sequence of nominal and real shocks and intertemporal propagation mechanisms. This takes away from the argument that current levels of employment (and low unemployment) are not sustainable over time. Karanassou, Sala and Salvador (2008) argue that different policies may be required to deal with shocks of differing durations and that chain reaction theory creates scope for applying policies to combat unemployment. For example, interventions designed to increase the capital stock and/or factor productivity could have persistent effects on the short- and long-run unemployment rate.

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Annex A1

TABLE A1.1
Description of the variables used

			Source
Endogenous variables			
l_t	Log active population[a]		INE
n_t	Log working population[a]		INE
u_t	Unemployment rate		$(l_t - n_t)$
w_t	Log real wage		INE
Exogenous and control variables			
k_t	Log capital stock[b]		IECON
prk_t	Log capital productivity[c]		BCU/IECON
z_t	Log working-age populationo[d]		INE
prn_t	Log labour productivity[e]		BCU/INE
Dummy variables			
d1	2002, third quarter	= 1 economic crisis	
d2	1997, first quarter	= 1 outlying value in series n_t and l_t	
d3	2000, second quarter	= 1 outlying value in series n_t and l_t	
d4	1994, fourth quarter	= 1 outlying value in series l_t	
d5	1987, third quarter	= 1 outlying value in series l_t	
d6	≥2002, third quarter	= 1 exit from crisis and economic recovery	
d7	1990, first to fourth quarter	= 1 high inflation, drop in real wage	
d8	1993, first quarter	= 1 outlying value in series w_t	
d9	1987, fourth quarter	= 1 outlying value in series w_t	

Source: National Institute of Statistics (INE), Central Bank of Uruguay (BCU) and Institute of Economics (IECON) of the University of the Republic, Uruguay.

[a] The series were constructed from the activity and employment rates given by the continuous household surveys and INE population projections.

[b] See Román and Willebald (2012) for a description of how this annual capital stock series is prepared. The data were converted into quarterly equivalents using a constant depreciation rate and quarterly investment information.

[c] This derives from the ratio between real gross domestic product (GDP) and capital stock.

[d] INE single-age population projections.

[e] This derives from the ratio between GDP and total hours worked.

<div align="center">

TABLE A1.2
F-test for the long-run relationship between variables

</div>

Labour demand						
$\Delta n_t = a_1 n_{t-1} + a_2 w_{t-4} + a_3 k_{t-1} + a_4 prk_{t-1} + \Sigma a_{ni}\Delta n_{t-i} + \Sigma a_{w4-i}\Delta w_{t-4-i} + \Sigma a_{ki}\Delta k_{t-i} + \Sigma a_{prki}\Delta prk_{t-i} + a_0$						
Lags		1	2	3	4	5
Selection criteria	Akaike	-5.55*	-5.49	-5.45	-5.41	-5.41
	Schwarz	-5.27*	-5.09	-4.96	-4.81	-4.70
Autocorrelation tests[a]	SC [χ^2 (1)]	0.51	0.83	0.75	0.64	0.17
	SC [χ^2 (4)]	0.83	0.86	0.65	0.40	0.19
F-test (H$_0$: $a_1=a_2=a_3=a_4=0$)[b]	Statistic	5.40	3.73	2.96	2.42	2.63
	Significance	**				
t-test (H0: $a_1=0$)[c]	Statistic	-3.52	-2.98	-3.13	-2.96	-2.72
	Significance	*				
Labour supply						
$\Delta l_t = b_1 l_{t-1} + b_2 w_{t-1} + b_3 u_{t-1} + b_4 z_{t-1} + \Sigma b_{li}\Delta l_{t-i} + \Sigma b_{wi}\Delta w_{t-i} + \Sigma b_{ui}\Delta u_{t-i} + \Sigma b_{zi}\Delta z_{t-i} + b_0$						
Lags		1	2	3	4	5
Selection criteria	Akaike	-5.83*	-5.79	-5.73	-5.67	-5.63
	Schwarz	-5.52*	-5.38	-5.22	-5.05	-4.91
Autocorrelation tests[a]	SC [χ^2 (1)]	0.97	0.64	0.39	0.54	0.94
	SC [χ^2 (4)]	0.56	0.60	0.30	0.46	0.59
F-test (H0: $b_1=b_2=b_3=b_4=0$)[b]	Statistic	5.25	3.55	3.20	2.33	2.57
	Significance	**				
t-test (H0: $b_1=0$)[c]	Statistic	-4.40	-3.41	-2.92	-2.28	-2.34
	Significance	***				
Wages						
$\Delta w_t = c_1 w_{t-1} + c_2 prn_{t-1} + c_3 u_{t-1} + \Sigma c_{wi}\Delta w_{t-i} + \Sigma c_{prni}\Delta prn_{t-i} + \Sigma c_{ui}\Delta u_{t-i} + c_0$						
Lags		1	2	3	4	5
Selection criteria	Akaike	-5.38	-5.41*	-5.40	-5.37	-5.38
	Schwarz	-5.10*	-5.05	-4.97	-4.86	-4.79
Autocorrelation tests[a]	SC [χ^2 (1)]	0.79	0.61	0.51	0.74	0.39
	SC [χ^2 (4)]	0.37	0.73	0.75	0.63	0.38
F-test (H0: $c_1=c_2=c_3=0$)[d]	Statistic	19.48	13.31	11.35	10.94	12.82
	Significance	***	***	***	***	***
t-test (H0: $c_1=0$)[e]	Statistic	-6.62	-5.73	-4.89	-4.65	-5.29
	Significance	***	***	***	***	***

Source: Prepared by the authors.

[a] *p*-value.

[b] Critical values for 90% (*): 2.72 - 3.77; for 95% (**): 3.23 - 4.35; for 99% (***): 4.29 - 5.61.

[c] Critical values for 90% (*): (-2.57) - (-3.46); for 95% (**): (-2.86) - (-3.78); for 99% (***): (-3.43) - (-4.37).

[d] Critical values for 90% (*): 3.17 - 4.14; for 95% (**): 3.79 - 4.85; for 99% (***): 5.15 - 6.36.

[e] Critical values for 90% (*): (-2.57) - (-3.21); for 95% (**): (-2.86) - (-3.53); for 99% (***): (-3.43) - (-4.10).

TABLE A1.3
System equation specification tests[a]

Estimations with instrumental variables[b]					
Labour demand		Labour supply		Wages	
NOR [χ^2 (2)]	0.925	NOR [χ^2 (2)]	0.845	NOR [χ^2 (2)]	0.408
SC [χ^2 (1)]	0.854	SC [χ^2 (1)]	0.087	SC [χ^2 (1)]	0.538
SC [χ^2 (4)]	0.788	SC [χ^2 (4)]	0.043	SC [χ^2 (4)]	0.543
HET [χ^2 (6)]	0.295	HET [χ^2 (8)]	0.148	HET [χ^2 (7)]	0.120

Source: Prepared by the authors.

[a] The values presented are the p-values for the normality tests (NOR), serial correlation with one and four lags (SC) and heteroskedasticity (HET).

[b] Instrumental variables: n_{t-1}, l_{t-1}, w_{t-1}, w_{t-4}, w_{t-5}, k_{t-1}, k_{t-2}, prk_{t-1}, prk_{t-2}, Δprk_t, Δprk_{t-1}, u_{t-1}, u_{t-2}, u_{t-3}, u_{t-4}, Δw_t, Δw_{t-1}, Δw_{t-2}, z_{t-1}, z_{t-2}, z_{t-3}, prn_{t-1}, prn_{t-2}, Δprn_{t-1}, Δprn_{t-2}, $d1$, $d2$, $d3$, $d4$, $d5$, $d6$, $d7$, $d8$, $d9$ and c.

Determinants of women's hours worked in Mexico: a pseudo-panel approach (2005-2010)

Armando Sánchez V., Verónica Villarespe R., Diego A. Román C. and Ana Liz Herrera M.

Abstract

The hours worked by Mexican women depend not only on wages and individual characteristics, but also on factors related to household structure, which generate incentives for women to restrict their hours of paid work. This study uses a pseudo-panel containing five million observations from the National Survey on Occupation and Employment, for 2005-2010. Different age cohorts of the female working population are analysed along with a pseudo-panel model that measures the sensitivity of women's hours worked to wage variations and factors related to household structure, such as the availability of help in the home and the presence of children. It is found that women's hours worked increase when the household contains another adult woman, but decrease in the presence of children or a male adult.

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I.   Introduction

Figures published by the Organization for Economic Cooperation and Development (OECD) show that Mexican women have been very slow to join the labour market in recent years, with their participation rate rising from 36.67% in 2005 to 38.32% in 2012. The empirical literature on various countries frequently attributes this trend to the dynamic of wages and individual characteristics. Nonetheless, no in-depth analysis has been made of the potential effects of other cultural, social and demographic factors on women's hours worked.[1] The central hypothesis of this study is that, aside from economic factors, household structure can also generate incentives for women to reduce their labour market participation. This is because women choose not only how many hours to allocate to paid work, but also the amount of time they will devote to producing household goods and care for family members —roles traditionally assigned women which compete with the time they spend in labour market activities (Acosta, Perticará and Ramos, 2006)—. For example, the presence of several small children in the household could create incentives for the woman to give up paid work and devote her time to care, owing to women's cultural and social roles, the lack of help in the household, and the shortage of public care services. In this situation, public policies need to be designed and implemented to free up part of the time that women spend on domestic and care activities in the home, to enable them to undertake paid work if they so wish. This generally means promoting policies that help reduce the social, cultural or demographic barriers that may restrict women's hours worked and their opportunities for advancement.

The target population of this study consists of women between 12 and 65 years of age. The determinants of hours worked were calculated using the pseudo-panel econometric technique, with a birth-cohorts approach. The database used includes about 5.2 million observations and contains information from the National Survey on Occupation and Employment (ENOE) spanning the third quarter of 2005 to the second quarter of 2010 (INEGI, n/d). In addition, elasticities were calculated for the sample using different estimation methods to ensure the statistical robustness of the results. The previous literature has no study of the effects of household composition on women's hours worked, based on such a large sample and using pseudo-panel econometric methods to increase the reliability of the estimated determinants, as was done in this study for Mexico (Deaton, 1997).

The results suggest that household structure affects the number of hours that women supply to the labour market. Firstly, there is more time available for paid work when there is another woman over 14 years of age in the household. This could represent unpaid help in domestic chores and household care, and would result in an increase in the number of hours available to supply to the labour market. Secondly, the presence of children has a negative effect on hours worked, because caring for them can put major demands on women's time (Arceo-Gómez and Campos-Vázquez, 2010). Consequently, household structure is an important determinant of the number of hours women have available for paid work. This could be explained in terms of the needs of household members constraining women's labour market decisions.

This article is divided into five sections including this Introduction. Section II makes a review of the existing literature and finds that previous research has suggested that household structure could affect women's hours worked. Section III puts forward empirical evidence at the cohort level on the average number of hours per week that women in Mexico spend on paid work. It also shows how hours worked vary with respect to household structure. Section IV sets out the results obtained from the pseudo-panel model and with different estimators, which guarantee the robustness of the results; and the fifth and final section offers a few final comments.

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[1]   Throughout this text, the term "hours worked" refers to the number of hours women devote to paid activities.

II. Previous evidence on female labour market participation and how it relates to household structure

The analysis of the factors that determine women's labour market participation has gained considerable importance in recent years. Domínguez and Brown (2013) make an in-depth study of the role of gender differences in labour market participation in Mexico. The key conclusions of that study stress that the presence of children and older adults in the household could affect a woman's decision to take on paid work, either from home or outside. The authors also suggest that paid work from home is viable in Mexico because care tasks, traditionally assigned to women, could make it difficult for women to participate in the formal labour market. Nonetheless, a decision to undertake paid work in the home could also reflect the shortage of schools and child-care centres with suitable hours to give mothers more options in terms of paid formal employment. In this connection, a study by Arceo-Gómez and Campos-Vázquez (2010) shows that the number of hours of paid work done by Mexican women with children under five years of age is much more sensitive to variations in the wage than those of the average Mexican woman. This could be explained by the fact that traditional female roles in the household put constraints on how women choose to distribute their time.

Attanasio, Low and Sánchez (2008) study the determinants of variations in the supply of labour by women with children, in three different cohorts in the United States, based on a life-cycle model. The results of their research show that a reduction in the monetary cost of child care and an increase in wages both encourage female labour market participation. In another study, Ludin, Mörk and Öckert (2008) estimate the effects of the reduction in child care costs on female labour supply in the United States and suggest that the impact is significant but heterogeneous, since it depends on the type of family and the region being studied. Warunsiri and McNown (2010) estimate the determinants of female labour supply in Thailand by constructing and analysing different synthetic cohorts of women. The authors' main finding is that there is a negative relation between the wage and women's hours worked in that country (the wage elasticity of labour supply was around -0.25). The article claims that the existence of a downward sloping labour supply curve could reflect to the competitive uses of a woman's time (Dessing, 2002). In other words, an increase in wages might reduce hours worked, because women could choose to devote more time to activities such as child care and the provision of domestic services.[2] Another important conclusion of the same study is that, although single women are more sensitive to wage changes than married women in terms of hours worked, they are less likely to be working. In this context, other research studies, such as those of Schultz (1990), Dessing (2002) and Warunsiri and McNown (2010) have concluded that the wage elasticity of hours worked is also negative in countries such as Thailand, Peru and the Philippines.

In other studies relating to developing countries in which pseudo-panel techniques are used, such as Bassi (2003), the conclusion is that, despite the wage elasticity of hours worked being positive as in developed countries, its magnitude is much smaller. Thus, women's hours worked in those countries might be explained by other variables that could be related to the use of time and the traditional tasks undertaken by women in the home, such as child care and domestic chores, such as washing, ironing and cooking. Frequently, one of the main motives for women entering the labour market is to maintain a certain level of household income, as indicated by Licona (2000) in a study on

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[2] For women who are working, a change in the wage induces both income and substitution effects, with opposing consequences for hours worked. Although the normal expectation is that the income effects will exceed the substitution effects, so that the elasticity of hours worked with respect to wages is positive, the evidence from several developing countries rejects that assumption.

the effect of poverty on female labour supply in Mexico. This has been corroborated in other studies, including Dasgupta and Goldar (2005) on indigenous women living in poverty. It can therefore be concluded that the determinants of female labour supply might be affected by very different factors than those affecting men (Juhn and Murphy, 1996).

Lastly, it is important to note that there are few empirical studies that focus on the effect of household structure on women's hours worked and, specifically, analyse whether the presence of other adults who could undertake unpaid domestic activities, or children in the home, would represent an assistance or a barrier to women's labour market participation. The present research aims to fill this gap in the traditional literature and provide empirical evidence using pseudo-panel techniques with a cohorts approach, applied to the relation between household structure and women's hours worked in Mexico.

III. Theoretical framework

The theoretical underpinning of the analysis in this article is the life-cycle model of labour supply of cohorts of women described by Attanasio, Low and Sánchez (2008). In that model, households face uncertainty about men's and women's wages, maternity is exogenous, and children impose a fixed monetary cost when the mothers decide to undertake a paid activity. Basically, the model explains changes in the female labour supply and assumes that households maximize their life-expectancy utility. The utility function that is developed is inter-temporally separable, and instantaneous utility depends on consumption per person in the home and the wife's labour supply choice.

The model assumes a household with an instantaneous utility function of the form:

$$u_t = u(c_t, \, P_t, \, e_t)$$

where P_t is a discrete variable {0,1} which measures the woman's labour supply decision; c_t is total household consumption and e_t is the number of equivalent adults in the household.

The model also defines $G(a_t)$ as the number of units of child care needed by a family whose first child is aged a_t. The price of each care unit is expressed as p. Thus, the total cost of child care incurred by a household when the woman participates in the labour market is given by:

$$F(a_t) = pG(a_t)$$

The study is relevant in terms of the woman's choice to join the labour market. The information provided shows that only the most productive women remain in the labour market after having children. It also shows the potential importance of the choice to continue working after having children, or not, and the repercussions of that decision on women's experience and wage.

IV. Data and results

1. Stylized facts on women's hours worked in Mexico

To analyse women's hours worked in Mexico, quarterly data were used from the ENOE, relating to the country's 32 states. The period of analysis runs from the third quarter of 2005 to the second

quarter of 2010. The ENOE is a nationally representative survey that generates occupational and sociodemographic statistical data. Although the sampling is random, the selected homes are replaced through a rotation scheme in which the 1/5 of the sample that has already completed a cycle of five visits from the questioners is replaced every three months. Thus, 80% of the sample is maintained each quarter.

The statistical analysis is applied only to women who fulfil the following characteristics: (i) they are employed; (ii) they undertook paid work during the previous week, and (iii) they have a monthly income. The sample includes women employed in both the private and public sectors, along with those who are self-employed. To provide empirical evidence of the relation between household structure variables and women's hours worked through time, a statistical analysis was performed at the cohort level. To that end, nine cohorts were constructed from the birth years of the target population, taking account of the characteristics mentioned above, and subject to the requirement of being of working age (from 12 to 65 years old).

The data show that the average number of hours women devote to paid work each week has trended upwards through the cohorts (see table 1). Women in cohort 1 (1940-1950) on average work 5.46 hours less than their peers in cohort 9 (1985-1992). Women in cohort 8 (1980-1985) apparently work most hours per week (40.33), but their wage level is not as high as cohort 4 (1960-1965) which on average works 2.9 hours less.

Table 1
Mexico: income and hours worked per week by cohort of employed women

Cohort		Average number of hours worked	Median real income[a]
1	1940-1950	34.55	1 718.37
2	1950-1955	36.36	2 237.03
3	1955-1960	36.99	2 582.30
4	1960-1965	37.40	2 767.25
5	1965-1970	37.87	2 724.89
6	1970-1975	38.15	2 691.84
7	1975-1980	38.92	2 724.89
8	1980-1985	40.33	2 643.89
9	1985-1992	40.01	2 148.98

Source: Prepared by the authors, on the basis of data from the National Survey on Occupation and Employment.

[a] Real income is defined as the monthly income in pesos declared by employed persons. To convert this into real values, the National Consumer Price Index (INPC) was used for each quarter in 2005-2010. The median wage is shown because the skewed distribution of the data makes it the best measure of central tendency.

The data in table 2 show that the presence of an adult in the home, which could represent unpaid help for women in domestic chores, is associated with an increase in hours worked. For example, the women of cohort 8 (1980-1985) on average work 4.6 hours more when there is another adult in the home than when there is not. In general, all of the cohorts benefit in terms of hours worked when there is another adult in the home. In fact, a women's willingness to enter the labour market seems to increase when there is potential assistance from some other adult in the household, so this could have a positive effect on the number of hours they devote to paid work.

Table 2
Mexico: average hours worked by women, by availability of help from another adult in the home, by cohort

Cohort		With help from another adult	Without help from another adult
1	1940-1950	34.99	33.85
2	1950-1955	36.33	36.44
3	1955-1960	36.96	37.10
4	1960-1965	37.46	37.19
5	1965-1970	38.20	37.17
6	1970-1975	39.34	36.66
7	1975-1980	41.02	36.45
8	1980-1985	41.73	37.12
9	1985-1992	40.46	37.28

Source: Prepared by the authors, on the basis of data from the National Survey on Occupation and Employment.
Note: A difference-in-means test was performed for the two groups (with assistance and without assistance) and the difference in hours worked proved statistically significant.

Table 3 shows that from cohort 3 onwards (1995-1960), the presence of at least one child under six years of age in the home is associated with fewer hours worked by women, compared to households that do not have children. This trend is maintained in the ensuing cohorts. Moreover, the number of hours by which women reduce their paid activity when they have children rises through the cohorts. In other words, women in cohorts 7 and 8 who have children under six years of age work 4.8 and 4.4 hours less, respectively, than those who do not have children. This means that caring for the children could have a negative effect on women's hours worked in Mexico, perhaps owing to the large amount of time needed. It was therefore decided to study the effect of this group of children since previous studies provide evidence that women with children under five participate less in the labour market than the average of women without children under five. This lack of association could be because children impose significant constraints on the allocation of time (Arceo-Gómez and Campos-Vázquez, 2010).

Table 3
Mexico: average hours worked by employed women, by presence of children under six years of age in the household, by cohort

Cohort		Have children under six	Do not have children under six
1	1940-1950	34.83	34.54
2	1950-1955	37.49	36.34
3	1955-1960	36.28	37.01
4	1960-1965	35.72	37.52
5	1965-1970	35.90	38.27
6	1970-1975	35.74	39.10
7	1975-1980	35.61	40.45
8	1980-1985	36.93	41.43
9	1985-1992	38.69	40.31

Source: Prepared by the authors, on the basis of data from the National Survey on Occupation and Employment.
Note: The difference in hours worked in both groups the cohorts 1, 2 and 3 is minimal because there are few observations of individuals in those cohorts with children under six. A difference-in-means test was performed for the two groups (with children under six and without them) and the difference in hours worked proved statistically significant.

The information shown in table 4 confirms that the presence of children in the household is strongly related to a reduction in women's hours of paid work. When children between six and 14 years of age are present, the average number of hours worked by women declines in all cohorts. In this case, the data show that women in cohort 7 (1975-1980) are those who work fewest hours when they have children in this age range. In contrast, the women in cohort 9 seem not to work less if they have young children, although that positive relation is present.

Table 4
Mexico: average hours worked by employed women, by presence of children between six and 14 years of age in the household, by cohort

Cohort		Have children from 6 to 14 years of age	Do not have children from 6 to 14 years of age
1	1940-1950	34.83	34.53
2	1950-1955	35.76	36.44
3	1955-1960	35.88	37.35
4	1960-1965	36.08	38.34
5	1965-1970	36.56	39.59
6	1970-1975	36.48	40.30
7	1975-1980	36.45	40.61
8	1980-1985	39.15	40.78
9	1985-1992	39.90	40.07

Source: Prepared by the authors, on the basis of data from the National Survey on Occupation and Employment.
Note: The difference in hours worked in both groups the cohorts 1, 2 and 3 is minimal because there are few observations of individuals in those cohorts with children under six. A difference of means test was performed for both groups (with children from 6 to 14 and without them) and the difference in hours worked proved statistically significant

The foregoing table suggests that the presence of an additional adult in the household is positively related to the number of hours that women can devote to paid work. In contrast, the presence of children in the household is inversely associated with the number of hours worked. The next section estimates models to test the statistical validity of those causal relations. Nonetheless, the situations in question could be reflecting the fact that children require care and attention, which could affect the availability of women's hours, and raise the cost of deciding to work if there is no help from another adult in the home.

Table 5 shows data for women who live in households containing at least one additional adult and at least one child aged under six. This reveals that in most of the cohorts, the presence of another adult who could do unpaid domestic chores in the household, enables women to spend more hours in paid work despite having young children. In this case, the largest difference between women with children under six who have help and those who do not, is in cohort 8 (1980-1985).

Another perspective on the relation between the presence of another adult and women's hours worked in the labour market is through their education levels. Table 6 shows that women with a low level of schooling work more when there is another adult in the home. In contrast, women with a high level of education do not seem to work more hours when there is another adult and small children in the home. It is important to note that women in cohorts 7 and 8 work the largest number of hours when they have help from another adult in the home, irrespective of their education level.

Table 5
Mexico: average hours worked by employed women who have children aged under six, by presence or absence of another adult in the household, by cohort

Cohort		With help from another adult and children under six	Without help from another adult and with children under six
1	1940-1950	35.01	33.86
2	1950-1955	37.75	36.45
3	1955-1960	36.92	37.17
4	1960-1965	36.31	37.47
5	1965-1970	36.20	37.77
6	1970-1975	36.80	37.77
7	1975-1980	38.65	38.40
8	1980-1985	42.10	39.96
9	1985-1992	41.53	40.28

Source: Prepared by the authors, on the basis of data from the National Survey on Occupation and Employment.
Note: A difference of means test was performed for the two groups (with assistance and without assistance) and the difference in hours worked proved statistically significant.

Table 6
Mexico: income and hours worked by employed women, by education level and cohort

Cohort		Preschool-primary		Secondary-upper middle		Higher		Postgraduate	
		Mean hours worked	Median real income[a]	Mean hours worked	Median real income[a]	Mean hours worked	Median real income[a]	Mean hours worked	Median real income[a]
With help from another adult									
1	1940-1950	35.07	1 595.98	36.48	3 040.62	35.53	5 901.05	34.52	9 116.52
2	1950-1955	35.99	1 833.22	37.51	3 133.63	36.22	6 347.52	37.13	12 301.48
3	1955-1960	37.11	1 917.40	37.63	3 133.63	35.87	6 319.28	37.83	13 584.27
4	1960-1965	37.38	1 922.74	38.48	2 982.71	35.69	5 830.01	38.62	9 349.95
5	1965-1970	37.47	1 949.60	39.24	2 706.95	36.51	5 410.39	36.99	9 837.72
6	1970-1975	38.01	1 960.05	40.53	2 588.95	38.35	5 008.08	35.05	8 232.13
7	1975-1980	40.31	1 983.68	42.34	2 566.51	39.49	4 637.43	39.51	5 549.22
8	1980-1985	42.75	2 066.33	42.90	2 472.20	39.09	4 099.05	38.26	5 224.20
9	1985-1992	41.59	1 854.15	40.56	2 133.64	37.50	3 159.64	26.00	4 435.16
Without help from another adult									
1	1940-1950	33.87	1 342.22	35.41	2 729.43	34.25	6 466.88	37.25	12 041.64
2	1950-1955	35.92	1 663.41	37.89	3 122.92	36.62	6 709.92	37.80	11 085.10
3	1955-1960	36.83	1 788.82	38.16	3 128.35	36.40	6 821.29	41.55	12 485.75
4	1960-1965	37.51	1 854.15	38.09	3 090.26	35.78	6 319.28	39.12	13 525.97
5	1965-1970	36.56	1 865.44	38.27	2 949.85	35.91	5 871.18	38.69	11 153.27
6	1970-1975	35.16	1 786.10	37.55	2 669.93	36.18	5 440.73	36.15	9 092.20
7	1975-1980	34.40	1 737.80	37.32	2 464.37	36.67	5 032.44	36.38	7 630.18
8	1980-1985	35.42	1 744.85	37.64	2 379.84	37.57	4 471.50	35.00	8 745.01
9	1985-1992	35.02	1 674.62	37.99	2 246.41	37.63	3 920.10	-	-

Source: Prepared by the authors, on the basis of data from the National Survey on Occupation and Employment.
Note: A difference of means test was performed for the two groups (with assistance and without assistance) and the difference in hours worked proved statistically significant.

[a]   Real income is defined as the monthly income in pesos declared by employed persons. To turn this into real values, the National Consumer Price Index (INPC) was used for each quarter in 2005-2010. The median wage is shown because the skewed distribution of the data makes it the best measure of central tendency.

The stylized facts developed in this section suggest the following: (i) the average number of hours worked by women in each cohort is greater when there is another adult in the household, and (ii) with the presence of children under 14 years of age, the average number of hours worked declines in each cohort. This provides empirical evidence that household structure probably affects the time women devote to paid work, which could be explained by the fact that women must also choose the number of hours they will spend looking after household members and producing goods in the household in an economy such as Mexico's. It was also found that education level can affect the number of hours that women are willing to spend in the labour market.

2. Results

To determine the effect of the household structure factors on women's hours worked, the following econometric specification was formulated:

$$ln\ h_{it} = \theta lnw + X_{it} + Z_{it} + u_i \tag{1.3}$$

where:
$ln\ h_{it} lnh$ is the logarithm of monthly hours worked by the women of cohort i in time t.
$ln\ w\ lnw$ is the logarithm of the monthly real wage.
$X_{it}\ X$ is a vector of socio-demographic variables and other individual characteristics.
$Z_{it}\ Z$ is a vector of household structure variables.
u_i is the error term.

Nonetheless, estimating the equation could raise the problem of endogeneity in the monthly real wage variable, owing to potential simultaneity between that variable and hours worked. To address this problem, instruments were constructed for the wage,[3] whose econometric specification includes the real exchange rate, the level of imports and the minimum wage (Robbins, Salinas and Manco, 2009).

Table 7 sets out the results of the model of the determinants of women's hours worked based on a pseudo-panel formed from the representative sample of the ENOE at national level. The results are shown with five different types of estimates to ensure their statistical robustness. The data shown in columns I to IV were estimated from a panel containing the whole sample of ENOE individuals, spanning the third quarter of 2005 to the second quarter of 2010, with about 5.2 million observations. Column (I) reports the estimation of the regression of hours worked without controlling for heterogeneity between individuals or wage endogeneity. Column (II) includes cohort and time effects. The coefficients shown in column (III) were calculated by correcting for the problem of wage endogeneity using instrumental variables. Column (IV) shows estimations based on weighted least squares (WLS). Lastly, column (V) shows the elasticity of women's hours worked, based on a dynamic pseudo-panel technique. In other words, after obtaining the mean of the variables of interest for each cohort, a temporary database is obtained, and lags are added to the dependent variable and instruments to control for possible endogeneity of the real wage.

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[3] The variables used as an instrument for the wage were: age, age squared, real exchange rate, imports and minimum wage.

Table 7
Mexico: factors determining women's hours worked

Variable	POOL				Dynamic HS pseudo-panel (V)
	(I)	(II)	Instrumental variables[a] (III)	Weighted least squares (IV)	
Wage	0.412 [0.0005]	0.411 [0.0005]	0.400 [0.0108]	0.412 [0.0005]	0.273 [0.1333]
Presence of another adult (woman) in the household	0.081 [0.0009]	0.081 [0.0009]	0.081 [0.0009]	0.081 [0.0009]	0.005 [0.0030]
Presence of another adult (man) in the household	-0.003 [0.0009]	-0.003 [0.0009]	-0.004 [0.0010]	-0.003 [0.0009]	-0.001 [0.0028]
Children aged 6-14 years	-0.027 [0.0005]	-0.027 [0.0005]	-0.028 [0.0007]	-0.027 [0.0005]	0.001 [0.0037]
Education	-0.032 [0.0001]	-0.032 [0.0001]	-0.031 [0.0010]	-0.032 [0.0001]	-0.029 [.0143]
Education$_{-1}$	-	-	-	-	0.010 [0.0132]
Education$_{-2}$	-	-	-	-	-0.020 [0.0164]
Age	-0.004 [0.0002]	-0.004 [0.0001]	-0.004 [0.0002]	-0.004 [0.0002]	0.091 [0.0386]
Age$_{-1}$	-	-	-	-	-0.187 [0.0409]
Constant	4.121 [0.0071]	4.093 [0 .0040]	4.139 [0.0200]	4.122 [0.0072]	0.0203 [0.0089]
Number of observations	1 486 014	1 486 014	1 441 978	1 486 014	261

Source: Prepared by the authors, on the basis of data from the National Survey on Occupation and Employment.
Note: All of the regressions include dummy regressors in each of the quarters. Regressions I, II and IV include cohort and time fixed effects, while II only includes cohort fixed effects.
[ ]: Standard deviation.

[a] Instrumentalized according to the methodology specified above.

The results show that there is a positive elasticity between the wage and women's hours worked, because a wage increase would make them more willing to increase the proportion of their time spent in paid work. The coefficients obtained in each of the models with different estimation methods are consistent in terms of their values and statistical significance. Column (V), which corresponds to the results of the dynamic pseudo-panel, indicates that with a 10% wage increase, women would be willing to work 2.7% longer. Nonetheless, according to the models presented, the wage is not the only factor with a positive effect on the hours women work in the labour market. Other factors related to household structures also affect the time women devote to paid work, which have not been studied in depth in the previous literature. For example, the results of the pseudo-panel show that when there is another woman in the household (over 14 years of age), who could represent extra unpaid help in household chores, women spend more time in paid work (0.005). This result is consistent with those of regressions I, II, III and IV, since the effects in all versions presented and in all cases are statistically significant. Based on these statistically robust results, it can be claimed that the presence of another woman in the household could imply help in the domestic chores that women normally fulfil, so the effect on the amount of time women spend in the labour market would be positive (the extra help will enable them to spend less time on household activities and taking care of household members). Public policy should therefore focus on providing help for women in the home, to reduce the time they have to spend on domestic chores, which often impedes their voluntary participation in the labour market. In other words, public policies are needed that reconcile or replace certain unpaid tasks, such as child care service or direct help for the care of household members who need it (Gammage and Orozco, 2008). Nonetheless, the presence of at least one male adult (over 14 years of age) in the household, without a paid job, produces a negative effect on women's hours worked (-0.001). Although this

coefficient is not statistically significant (only in the pseudo-panel regression), it displays a robust trend in all models reported.

Regressions I to IV show that the effect on hours worked of the presence of children between six and 14 years of age is negative and statistically significant. Thus, when the household contains children, women reduce the time they spend on paid work. This can be explained by the time demands that child care represents for women (Pedrero, 2009). This suggests that a reduction in childcare costs increases women's hours worked in the labour market (Attanasio, Low and Sánchez, 2008). Another relevant and consistent finding in all of the estimations is the negative impact of education level on the hours women spend in paid work. For that purpose, the variable that indicates women's years of schooling was considered; and, as higher levels of schooling are associated with higher incomes and better job conditions (Domínguez and Brown, 2013), the substitution effect might dominate in this case. In other words, if they already have a substantial level of income, women could decide to reduce the number of hours worked in the labour market and spend them on other activities. Another factor that has a negative effect on women's hours worked in the labour market is age, since the result is also robust in all versions presented in table 7. On this basis, it can be claimed that the older the women, the fewer hours they are willing to spend in paid work.

In general, the variables that positively affect women's hours worked and whose results display an adequate level of statistical robustness are: (i) the wage, and (ii) the presence of another woman over 14 years of age in the home. Factors that have a negative effect on women's hours worked and are consistent in all regressions are: (i) the presence of a male older than 14; (ii) the presence of children aged between six and 14; (iii) education level, and (iv) age. On this basis, it can be stated that household structure is an important element in determining the number of hours women spend in paid work, and that this could be related to the time they devote to household activities.

V.   Final comments

The aim of this study has been to determine whether household structure in any way affects the number of hours that women spend doing paid work in the labour market. The results obtained with the pseudo-panel technique suggest that a woman's hours worked do depend on the number of adults and children living in the household. Specifically, evidence was found that the presence of young children reduces the availability of women's time. This could be directly related to the care and attention that children require of the woman, based on her traditional role within the household, thereby occupying most of her time and simultaneously reducing her availability for working outside. In this context, it can be stated that women undertake a considerable amount of unpaid work in the household, and this restricts their possibilities of entering the labour market. The results of this study admit several public policy suggestions that focus on reducing the time women spend on domestic activities and care in the home. A first recommendation is to design and develop actions to generate and provide incentives for help in the home. Although the presence of another woman makes it possible to reduce the workload in the household and promotes an increase in female hours worked, public policies are needed to replace that help, so as to prevent unpaid work being transmitted from one woman to another. Accordingly, public policies are needed —such as an increase in the number of public child-care centres and the development of school programmes— that give greater flexibility to women's timetable. It is important to develop public policies aimed at supporting domestic activities, particularly those related to household members and their care. With this type of action, women's hours worked would increase, because they would be better placed to supply a larger number of hours to the labour market. This would also enable them to improve their quality of life should they

voluntarily so decide. Domestic activities, such as caring for family members and producing of goods in the household, represent one of the key factors that restrict the time women have available for formal work; and this affects women much than men (INEGI, 2012).

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The intersection between class and gender and its impact on the quality of employment in Chile[1]

Omar Aguilar, Pablo Pérez, Rubén Ananías, Claudia Mora and Osvaldo Blanco

Abstract

This study explores the impact of the intersection between class and gender on the quality of employment in Chile. The method used to measure social class position is based on the work of Erik O. Wright, while, for the quality of employment, a multidimensional measurement was used, including one index for objective working conditions and two indices for subjective ones (motivation on the job and the perception of control over work processes). The results demonstrate that class and gender give rise to significant differences in objective and subjective job quality. However, the data also indicate that gender (more specifically, the fact of being female) does not necessarily amplify the class-based inequalities observed in the labour market. Drawing on these findings, a number of thoughts about how the class/gender intersection operates in the Chilean labour market are shared in the final section of this study.

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I.   Introduction

One approach to the analysis of inequality is based on the concept of intersectionality, which is also known as "intersectional discrimination." In this approach, it is assumed that economic and social inequalities are not solely a consequence of a person's position within the production structure (i.e. a person's class position) but are also a result of other factors and categories, such as gender, according to which people can be classified. The concept of intersectionality can therefore be used to gain an understanding of the processes by which class and gender, taken together, give rise to differentiated access to opportunities and resources.

This study focuses on measuring the intersectionality of class and gender and on gauging how these two factors influence job quality. Using the class-based scheme proposed by Erik O. Wright (1994) as a point of departure, an effort is made to quantify the ways in which class, gender and the intersection of these two factors translate into variations in job quality, with the latter being analysed on the basis of multidimensional measurements of objective and subjective employment and working conditions.

The aim of this study is to contribute to the development of multidimensional measurements of job quality in Chile and to the analysis of class and gender by examining the extent to which the intersection of these two factors provides a means of gaining a better understanding of the inequalities existing in highly flexible labour markets such as that of Chile. To that end, the study is divided into seven sections. Following this Introduction, section II presents an analysis of the literature dealing with class-based models as viewed from a gender perspective. Section III offers a discussion of the measurements of multidimensionality that have been applied to job quality, while section IV outlines the hypotheses on which this study is based. Section V reviews the data, variables and methods used in the research. Section VI presents and discusses the research results. Section VII provides an overview of the main findings of this study.

II.   Class-based models and a gender perspective

The most commonly used approach to the study of social inequality is based on an analysis of the positions that people occupy in the social structure. The concept of class has thus been frequently used in sociological studies that seek to determine the positions occupied by different individuals in production and market processes and to explore how that position affects their levels of material well-being and their life opportunities (Crompton, 2008).

Class position is not the only determinant of people's life opportunities, however. There are other dimensions in today's societies that also influence these outcomes. Sex, or gender, is one example (Crompton, 1989).[2] The available empirical evidence shows that gender is a core determinant of the opportunities that are open to people in the labour market (Browne and Misra, 2005; Stier and Yaish, 2014).

This has led researchers to look more closely into the relationships between class and gender. In the 1960s, the feminist movement engaged in a debate concerning the theoretical and methodological implications of the analysis of women's positions in the social structure (Pollert, 1996; Ferree and Hall, 1996; Yuval-Davis, 2006; Davis, 2008). As more and more women entered the labour market, they

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[2]   In the English-language literature, the term "gender", rather than "sex," is generally used to refer to the natural status of a person, as distinct from the social construct that may be associated with it. In this study, the term "gender" will, in most cases, be used to refer to a person's sex as viewed from the standpoint of the relationship it bears to the availability of resources and opportunities.

began to question the characterization of women as a peripheral component of the class system, which was, according to this point of view, reflected in the fact that class position was analysed on the basis of the occupational status of the head of household and chief breadwinner, who was generally a man.

The large-scale entry of women into forms of gainful employment in advanced capitalist societies prompted researchers to ask themselves to what extent sex was independent of class. They discussed, for example, how to go about analysing situations in which there were two heads of household who occupied different class positions. In the realm of empirical research, this debate raised questions as to which unit of analysis was appropriate, i.e. whether it was better to gather data at the individual or household level (Baxter, 1992).

The most well-known stance regarding the central importance of the household in studies of social class is that of John Goldthorpe (1983). In his view, all members of the household occupied the same class position. He argued that class position should be measured on the basis of the economic activity conducted by the man of the house because men were the main providers and breadwinners.

In contrast, feminists maintained that, given the fact that there were some households that were economically dependent on a woman and there were some in which both the man and the woman were breadwinners, it was necessary to have a joint classification model, i.e. a model that was capable of combining the attributes of both spouses in determining their class or status (Baxter, 1992). These discussions led to the development of an approach based on the concept of intersectionality, which focuses on the ways in which the interactions of various dimensions of inequality influence life opportunities.

Not all schools of thought have embraced the concept of intersectionality, however, with one of the main reasons for this being that it entails a complete overhaul of the theoretical assumptions underlying the way in which empirical data are interpreted. Be that as it may, Wright (1989 and 1997) has conducted empirical research on class and gender in which he demonstrates that gender is an extremely important determinant of access to positions of authority in countries such as Australia, Japan, Sweden and the United States.

However, in an effort to vindicate the Marxist theory of social class, Wright (1992) states that social class is a "gender-neutral" abstract concept in much the same way as patriarchy is, in the abstract, a "class-neutral" concept. In other words, in the abstract, class and gender can be understood as two totally distinct concepts. Accordingly, Wright contends that the complex relationship between class and sex can only be understood, in the abstract, if they are thought of as independent phenomena.

Based on this line of reasoning, Wright contends that the interaction between class and gender exists, but only at a concrete level. In other words, class structures are shaped by gender relations solely in a circumstantial, material sense. By the same token, it is only at that concrete, circumstantial level that gender shapes other class-related phenomena, such as class consciousness and collective action (Wright, 1992, p. 47).

In this study we follow Wright's line of thinking and define intersectionality between class and gender as a concrete, rather than as an abstract, phenomenon. On that basis, we explore how intersectionality is manifested, concretely and systematically, in the Chilean labour market.

Few quantitative studies on intersectionality have been conducted (Lovell, 2006), and those that do exist have used a wide variety of methodologies. While some are based on a comparison of wage gaps or the average wages of men and women (Browne and Misra, 2005; McCall, 2001 and 2005), others employ correspondence analysis, cluster analysis or discriminant analysis (Andes, 1992; Jaoul-Grammare, 2007). There have also been studies that examine the relationship between class and gender by looking at how gender influences the distribution of individuals in the class

structure. For example, using logit regression models, Mjøset and Petersen (1983) have demonstrated that women are much less likely to occupy class positions of authority (e.g. supervisory positions) than men.

In this study, a regression analysis will be used to examine the effect of class/gender intersectionality on the labour market using multidimensional measurements of objective and subjective job quality. The focus is on the labour market because it is one of the main areas in which the lasting effects of class and gender are evident (Stier and Yaish, 2014; Armstrong, Walby and Strid, 2009). Both in Chile and in other capitalist societies, the labour market is the main conduit for the allocation of resources and remuneration to individuals.

III. Multidimensional measurements of job quality

For some years now, scholars have shown interest in constructing job quality measurements that incorporate the dimensions of employment conditions, working conditions and the social/workplace environment (Burchell and others, 2012). Measurements of this type have been proposed in various countries and institutions, including Canada (Statistical Institute of Quebec, 2008), the United Kingdom (Green, 2006), the United States (Handel, 2005) and France (Guergoat-Larivière and Marchand, 2012; Ralle, 2006). In Europe, both the European Commission and the European Foundation for the Improvement of Living and Working Conditions have made valuable contributions in this direction (Leschke, Watt and Finn, 2008 and 2012; Green and Mostafa, 2012; Eurofound, 2012). The Organization for Economic Cooperation and Development (OECD) and the International Labour Organization (ILO) have also been moving forward with a line of work that involves the development of multidimensional employment quality metrics (Bescond, Chataignier and Mehran, 2003; Anker and others, 2003; Davoine, Erhel and Guergoat-Larivière, 2008). For the most part, these proposed systems of measurement include such factors as income, job stability, working time, training and worker autonomy, among other dimensions.

Ruiz-Tagle and Sehnbruch (2011) have proposed a multidimensional measurement of job quality in Chile that includes four dimensions: income level, type of contract and access to social security coverage, job seniority and training. The proposal that will be put forward here differs from that model in that it also includes subjective aspects, such as workers' own perceptions of their working conditions and their autonomy in the workplace.

The authors of recent research papers on the subject do not agree as to how to go about creating a synthetic measurement of employment quality. Our choice of indicators for measuring objective and subjective aspects of job quality was guided by the suggestions made in the specialized literature on the subject and by the results of a number of statistical tests that we ran (in particular, Cronbach's alpha coefficient and principal components analyses).

One problem with measuring the quality of employment is that, apart from income, none of the other measures in the models discussed in the literature incorporates a universal unit or metric; for the most part, nominal and/or ordinal variables are used. What the literature recommends in this connection is that indices are used only insofar as the procedures used for their construction and interpretation are transparent (Leschke, Watt and Finn, 2008; Green and Mostafa, 2012).

There are no universally accepted rules regarding how to weight the different indicators either. The different approaches used to determine the relative weighting of each indicator range from purely normative models to ones that rely entirely on empirical criteria. The general rule, however, is for most of the indices to be simply the sum of each of their components, with the result that each

component (variable) of a given index will have an equal relative weight (Guergoat-Larivière and Marchand, 2012).

These general recommendations were followed when building the indices for this study. However, the multidimensional measures of employment quality used here differ from the models reviewed in the literature in that they include subjective variables —that is, variables that measure workers' own perceptions of their working conditions. This study therefore is also intended to contribute to the debate on how to construct better measurements for use in empirical studies of job quality.

IV. Hypothesis

In this study, an effort will be made to compare two different hypotheses based on the results of studies on class, gender and the interaction between the two (Andes, 1992; Anthias, 2001; Browne and Misra, 2005; Wright, 1997; Mintz and Krymkowski, 2010).

The first (hypothesis 1) is a general one according to which gender and class have a significant impact on employment quality such that people in the working class (or in positions proximate to that class) and women are likely to have lower-quality jobs. More specifically, this hypothesis posits that women and members of the working class (or those in proximate positions) will have lower scores on the objective and subjective indices of employment quality.

The second hypothesis (hypothesis 2) is more specific than the first and is drawn from the findings of analyses of class/gender intersectionality. According to this line of thinking, the impact of social class on employment quality may be different for men than for women. More specifically, the difference may stem from the fact of being female could amplify class differences (especially those associated with membership in a subordinate class). Thus, for example, the quality of employment would be lower for members of the working class than for highly skilled professionals and would be even lower for working-class women than for working-class men.

In sum, based on these two hypotheses, the expectation is that working-class individuals will score lower on the objective and subjective indices of job quality (hypothesis 1) and that working-class women will score even lower (hypothesis 2).

V. Data, variables and methods

This study draws on data from the National Survey on Employment, Work, Health and Quality of Life of Workers in Chile conducted in 2009 and 2010 by the Ministry of Health and the Ministry of Labour and Social Security of Chile. The survey sample and its selection (9,503 cases chosen using a probability sampling method) are such that the data are representative of all Chilean workers over the age of 15 who were employed at the time that the survey was carried out or that, even if they were unemployed at that time, had had a job in the formal or informal sector within the preceding 12 months.

This database was chosen because it includes a series of variables that can be used to measure the relationship between gender and class and their impact on job quality. Standard multivariate analysis techniques —specifically, linear regression models based on ordinary least squares estimates— were used. These techniques were the most appropriate not only because the interpretation of the coefficients is straightforward, but also because all the dependent variables are interval variables (specifically, indices that were constructed to represent the objective and subjective components of job quality).

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1. Dependent variables

The main dependent variable in this study is "job quality" as measured by three indices: an objective job-quality index and two subjective job-quality indices.

(a) The objective job-quality index

In line with earlier studies (Leschke, Watt and Finn, 2008 and 2012; Davoine, Erhel and Guergoat-Larivière, 2008; Green and Mostafa, 2012; Ruiz-Tagle and Sehnbruch, 2011), this index was built on the basis of two variables.

The first, "job security," was constructed by cross-tabulating the following variables: (i) the type of employment (temporary or permanent), and (ii) social security coverage. This last variable was generated by cross-referencing two questions: "Are you enrolled in a social insurance system?" and "At this point in time, are you making payments into that system or is your employer doing so on your behalf?". The responses to these two questions were combined to generate a variable that could take any one of three possible values: 0 = is not a member of a social insurance system and social insurance contributions are not being made; 1 = is a member but social insurance contributions are not currently being made; and 2 = is a member and social insurance contributions are being made.

Thus, the variable for job security is the result of a cross-tabulation of the type of employment and of social security coverage and contributions (or the absence thereof). The outcome was a variable whose values may range from 0 (does not have social insurance coverage and is employed on a temporary basis) and 5 (does have social insurance coverage, contributions to that system are being made and has a permanent job). The values between those two extremes represent various intermediate situations (e.g. has social insurance coverage but is employed on a temporary basis or does not have social insurance coverage but has a permanent job). The higher the score for this variable, the better the quality of employment is.

The second variable was "income level." This variable was constructed on the basis of the standard question regarding the income of survey respondents. Their answers were recorded using values ranging from 0 to 5 (the higher the score, the higher the income level).

The values for these two variables were added together in order to obtain the final index for objective job quality, with higher scores indicating a greater degree of objective job quality.

(b) The subjective job-quality indices

Subjective perceptions of job quality were measured on the basis of two specific dimensions: motivation on the job and the perception of control over work processes.

(i) Motivation on the job

The motivation index was constructed on the basis of the simple sum of responses to three questions on a Likert scale. The questions were: "Do you feel that the work you do is important?", "Do you feel motivated in your job and committed to your work?" and "Do you enjoy the work you do?". The values for the responses to each of these three questions ranged from 1 (never) to 5 (always).

(ii) Perception of control over work processes

The index of perception of control over work processes was constructed on the basis of the simple sum of responses to four questions on a Likert scale. These questions were: "Are you

able to influence the amount of work that is assigned to you or that you have?", "Are you able to choose or change the order in which you do your various tasks?", "Are you able to choose or change the way in which you do your work?" and "Are you able to decide when to take a break?". The responses to each of these questions could range from 1 (never) to 5 (always).

In order to make the three indices of job quality comparable, the objective index and the two subjective indices were standardized on a scale of from 0 to 100, with a higher score corresponding to better job quality or a perception of better job quality. As shown in table 1, Cronbach's alpha and the principal component analyses indicate that the indices are one-dimensional and internally consistent. Cronbach's alpha for the objective quality index is lower than it ought to be (0.52), but it was retained nonetheless because it includes a number of elements (job stability, social insurance coverage and income) that are commonly used to measure job quality. The components of this index are also particularly relevant for an analysis of the relationship among class, gender and job quality. Unlike some of the measurements that are generally used for this purpose (e.g. type of employment contract), these items are applicable both to wage earners and to independent workers (entrepreneurs or the self-employed). Even though this objective quality index is one-dimensional, its low Cronbach's alpha coefficient suggests that the results of the analysis should be interpreted with caution.

Table 1
Employment quality indices

Indices	Variables	Mean	Standard deviation	Minimum value	Maximum value
Objective job quality (n = 9.248) Cronbach's alpha: 0.52 Eigenvalue, factor 1:1.37	1. Level of job security: type of employment (permanent or temporary) + social security coverage 2. Income level	50.82	23.37	0	100
Motivation (n = 9.177) Cronbach's alpha: 0.82 Eigenvalue, factor 1:2.20	1. Feels that the work he or she performs is important 2. Is motivated and committed to his or her work 3. Enjoys his or her work	84.72	21.07	0	100
Perception of control over work processes (n = 9.111) Cronbach's alpha: 0.83 Eigenvalue, factor 1:2.68	1. Can influence the amount of work assigned to him or her 2. Can change the order in which tasks are performed 3. Can change the way in which he or she does the work 4. Can decide when to take a break	55.97	32.63	0	100

Source: Prepared by the authors, on the basis of data from the National Survey on Employment, Work, Health and Quality of Life of Workers in Chile (ENETS, 2009-2010).

2. Independent variables

The independent variables are gender and social class. Gender was included as a dummy variable, with the reference category being "male."

The variable of social class was defined on the basis of the work of Wright (1994). According to Wright's analytical scheme, there are three variables that determine the main class positions in today's capitalist societies: (i) the ownership of means of production or the lack thereof, from which a distinction is drawn between wage earners and independent workers; (ii) the possession of authority over work processes or the lack thereof, and (iii) the possession of skills or the lack thereof. Wright uses these last two criteria to distinguish between different positions within the wage-earning population.

Based on these general criteria, the variable "employment situation" ("as your main occupation, you work as…") was then used to distinguish between those who own means of production and those who do not. Owners of means of production are classified as big employers (business owners with 10 employees or more), small employers (business owners with between 2 and 9 employees) and self-employed persons.

The wage-earning population is divided into subcategories based on the authority over work processes and the skill level. A series of variables are used to measure the extent of a survey respondent's authority (e.g. if he or she supervises the work of others, is authorized to hire or dismiss employees, can oversee and organize the work of others or occupies a management, supervisory or subordinate position in the workplace). Based on these variables, people are classified as managers, supervisors or workers.

A person's skill level is gauged on the basis of the International Standard Classification of Occupations (ISCO-88), which classifies occupations —especially in the case of those in the upper echelons— primarily in terms of the qualifications and skills required to carry out a given task or job (Ganzeboom and Treiman, 2003). Based on the 27 ISCO occupations listed at the two-digit level, survey respondents can be classified as "experts" (highly qualified professionals), "semi-skilled workers" (technicians and associate professionals) and "unskilled workers" (occupations for which no formal qualifications are required).

One major modification has been made in Wright's scheme: the addition of the category of "informal petty bourgeoisie." Since, in Latin America, many self-employed persons and small business owners conduct their economic activities in the informal sector (Tokman, 2009; Ministry of Economic Affairs, 2013), this class category was added in order to distinguish between those people and persons who, although they belong to the petty bourgeoisie or are small-scale employers, are linked to the formal sector of the economy. While there are many different definitions of informal labour and the informal sector (Portes and Haller, 2004), given the type of data that were available, we decided to use the general criteria proposed by the Regional Employment Programme for Latin America and the Caribbean (PREALC, 1978). Thus, "informal petty bourgeoisie" is defined as the category composed of self-employed persons and small-scale employers who are engaged in unskilled or low-skilled activities (i.e. those whose occupations are in ISCO-88 Major Groups 5 through 9). Based on these criteria, we constructed the classification of 13 social classes that is shown in table 2.

Table 2
Distribution of employed persons, by class position (13 classes) and gender
(Frequency and percentages)

	Men		Women		Total	
	Frequency	Percentage	Frequency	Percentage	Frequency	Percentage
1. Bourgeoisie	26	0.4	7	0.2	33	0.4
2. Small-scale employers	99	1.7	42	1.2	141	1.5
3. Formal petty bourgeoisie	106	1.8	65	1.9	171	1.9
4. Informal petty bourgeoisie	1 265	21.9	578	17.0	1 843	20.1
5. Expert managers	23	0.4	11	0.3	34	0.4
6. Expert supervisors	112	1.9	96	2.8	208	2.3
7. Non-managerial experts	85	1.5	171	5.0	256	2.8
8. Skilled managers	35	0.6	12	0.4	47	0.5
9. Skilled supervisors	285	4.9	121	3.6	406	4.4
10. Skilled workers	814	14.1	585	17.2	1 399	15.2
11. Unskilled managers	39	0.7	11	0.3	50	0.5
12. Unskilled supervisors	339	5.9	159	4.7	498	5.4
13. Unskilled workers	2 554	44.2	1 547	45.4	4 101	44.6
Total	5 782	100	3 405	100	9 187	100

Source: Prepared by the authors, on the basis of data from the National Survey on Employment, Work, Health and Quality of Life of Workers in Chile (ENETS 2009-2010).

Since the main objective of this study is to analyse how the intersection of class and gender influences job quality, the categories in the 13-class scheme in which there were very few cases were merged. Thus, the categories "bourgeoisie" and "formal petty bourgeoisie" were combined to form the category of "small-scale employers." For the same reason, "expert managers," "expert supervisors," "skilled managers" and "skilled supervisors" were merged into a single category. These two modifications —grouping all employers into a single category and all persons whose skill level qualifies them as experts with managerial functions in another— resulted in a scheme composed of nine class positions. This was the scheme used for running the regressions.

The demographic controls and economic variables described below were also included in the regression models.

(a) Demographic controls

Three demographic controls, apart from class and gender, were included in the regression models: age (in years); level of education (measured using four dummy variables, with the reference category being "completed primary education or less") and place of residence (1 for "Metropolitan Region" and 0 for "other region").

(b) Economic variables

Two economic variables were also included in the regression models: "economic sector" (agriculture, manufacturing or services, with "services" being the reference category) and public sector or private sector (with "public sector" being the reference category).

A number of regression models (ordinary least squares estimates) were used to see how the intersection of class and gender translates into variations in job quality. The variables were introduced in the same order in all cases, with the main independent variables (social class and gender) being incorporated before the demographic controls and economic variables.

VI. Results

1. The effect of class and gender on job quality

The figures shown in table 3 indicate that, together with members of the working class, the formal and informal petty bourgeoisie have the poorest-quality jobs. The size and similarity of the coefficients for these latter two groups —which are even higher than the coefficient for the working class— demonstrate the substandard nature of job quality among self-employed persons. This is true even in those cases where the self-employed persons' skill levels are such that it can be assumed that they are working in the formal sector (formal petty bourgeoisie).

In all of these cases, the correlation between being a woman or a member of the subordinate class (or both) and having a lower-quality job holds even after introducing the demographic controls and economic variables in models 2 and 3.

Table 3
Coefficients of objective job quality determinants in Chile

	Model 1	Model 2	Model 3
Gender (reference: male)			
Female	-9.233***	-10.24***	-9.897***
	(0.442)	(0.427)	(0.460)
Social class (reference: small-scale employers)			
2. Formal petty bourgeoisie	-20.00***	-18.09***	-18.21***
	(2.247)	(2.158)	(2.141)
3. Informal petty bourgeoisie	-29.53***	-21.03***	-20.95***
	(1.673)	(1.644)	(1.631)
4. Expert managers	8.804***	8.694***	9.267***
	(1.780)	(1.712)	(1.705)
5. Non-managerial experts	13.30***	6.132**	6.519**
	(2.55)	(2.066)	(2.052)
6. Skilled workers	-8.940***	-2.490	-2.055
	(1.694)	(1.653)	(1.642)
7. Unskilled managers	-5.011	1.021	1.192
	(3.353)	(3.222)	(3.198)
8. Unskilled supervisors	-0.592	5.093***	4.829**
	(1.838)	(1.789)	(1.778)
9. Unskilled workers	-14.86***	-6.338***	-6.096***
	(1.635)	(1.610)	(1.602)
Demographic variables			
Age		0.132***	0.134***
		(0.0166)	(0.0166)
Secondary education		10.83***	9.500***
		(0.490)	(0.501)
Vocational institute/technical training centre		15.70***	14.13***
		(0.808)	(0.817)
University or higher		24.63***	23.15***
		(1.146)	(1.147)
Residence in the Metropolitan Region		2.560***	1.861***
		(0.527)	(0.527)
Economic variables			
Agricultural sector			-5.896***
			(0.619)
Manufacturing sector			1.675**
			(0.537)
Private sector			2.978***
			(0.698)
Constant	67.65***	46.48***	45.04***
	(1.609)	(1.866)	(2.014)
Adjusted R^2	0.26	0.32	0.33
N	8 894	8 868	8 868

Source: Prepared by the authors, on the basis of data from the National Survey on Employment, Work, Health and Quality of Life of Workers in Chile (ENETS 2009-2010).

Note: Unstandardized coefficients, ordinary least squares regression.
The omitted variables are male (for gender), employers (for social class), primary education (for level of education), does not live in the Metropolitan Region (for place of residence), services sector (for economic sector) and public sector (for public or private sector).
Standard errors are shown in brackets. *** $p<0.001$; ** $p<0.01$; * $p<0.05$.

Table 4 gives the determinants for the two subjective job-quality indices. In the first index (motivation on the job), it can be seen that gender and class have a significant impact. As was true in the preceding case, the condition of being a woman and/or belonging to a subordinate class —e.g. the informal petty bourgeoisie or the skilled or unskilled working class— is associated with lower levels of motivation. In both cases, this correlation remains robust even after the demographic controls and economic variables are introduced.

The results are different, however, when the determinants of the second subjective job quality index are analysed (perception of having control over work processes). Contrary to the initial hypothesis, the condition of being a woman does not give rise to any significant difference in that perception. In the case of social class, two factors are noteworthy. First, none of the classes composed of persons who own means of production —not even the lowest-ranking informal petty bourgeoisie— is associated with any significant decrease in the perception of control over work processes. This suggests that ownership of means of production —no matter how modest those means may be— is a highly important consideration in understanding how people perceive their jobs. Second, and in contradistinction to the foregoing, the data indicate that the impact of membership in a subordinate wage-earning class is quite large. Thus, for example, membership in the skilled or unskilled working class translates into a decrease of nearly 35 points on the index of perceived control over work processes.

Table 4
Coefficients of determinants of subjective job quality in Chile

	Motivation on the job			Perception of control over work processes		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
Gender (reference: *male*)						
Female	-3.545***	-3.132***	-2.838***	0.173	0.573	-1.231
	(0.451)	(0.450)	(0.488)	(0.643)	(0.645)	(0.697)
Social class (reference: *small-scale employers*)						
2. Formal petty bourgeoisie	-2.165	-1.829	-1.674	0.586	1.165	1.262
	(2.253)	(2.232)	(2.230)	(3.252)	(3.246)	(3.233)
3. Informal petty bourgeoisie	-6.041***	-5.407**	-5.449**	1.915	2.464	2.761
	(1.661)	(1.686)	(1.685)	(2.410)	(2.461)	(2.451)
4. Expert managers	-1.517	-1.050	-1.472	-16.62***	-15.99***	-17.62***
	(1.768)	(1.753)	(1.757)	(2.561)	(2.561)	(2.560)
5. Non-managerial experts	-1.003	-1.901	-2.089	-28.38***	-29.65***	-31.01***
	(2.054)	(2.126)	(2.126)	(2.965)	(3.098)	(3.089)
6. Skilled workers	-8.837***	-6.958***	-7.152***	-34.78***	-32.87***	-33.85***
	(1.682)	(1.696)	(1.695)	(2.437)	(2.474)	(2.466)
7. Unskilled managers	-3.971	-2.474	-2.875	2.340	3.942	2.658
	(3.311)	(3.288)	(3.286)	(4.767)	(4.763)	(4.746)
8. Unskilled supervisors	-7.640***	-5.706**	-6.234**	-21.39***	-19.23***	-19.60***
	(1.836)	(1.846)	(1.847)	(2.656)	(2.689)	(2.681)
9. Unskilled workers	-14.16***	-12.16***	-12.63***	-35.18***	-33.13***	-34.02***
	(1.621)	(1.650)	(1.652)	(2.352)	(2.410)	(2.406)

Table 4 (concluded)

	Motivation on the job			Perception of control over work processes		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
Demographic variables						
Age		0.168***	0.165***		0.167***	0.144***
		(0.0175)	(0.0176)		(0.0251)	(0.0251)
Secondary education		0.938†	0.727		-0.352	-0.0739
		(0.518)	(0.534)		(0.742)	(0.763)
Vocational institute / technical training centre		2.178*	1.986*		1.910	2.092†
		(0.846)	(0.861)		(1.213)	(1.232)
University or higher		3.483**	3.299**		3.167†	3.195†
		(1.192)	(1.202)		(1.717)	(1.726)
Residence in the Metropolitan Region		-5.430***	-5.569***		-4.438***	-4.204***
		(0.542)	(0.545)		(0.777)	(0.779)
Economic variables						
Agricultural sector			-0.670			0.312
			(0.658)			(0.941)
Manufacturing sector			2.440***			-1.484†
			(0.567)			(0.811)
Private sector			-1.449*			-8.735***
			(0.738)			(1.057)
Constant	95.68***	86.93***	88.24***	80.05***	71.87***	82.08***
	(1.593)	(1.930)	(2.104)	(2.314)	(2.809)	(3.044)
Adjusted R^2	0.06	0.08	0.08	0.22	0.23	0.24
N	8 886	8 860	8 860	8 822	8 795	8 795

Source: Prepared by the authors, on the basis of data from the National Survey on Employment, Work, Health and Quality of Life of Workers in Chile (ENETS 2009-2010).
Note: Unstandardized coefficients, ordinary least squares regression.
The omitted variables are male (for gender), employers (for social class), primary education (for level of education), does not live in the Metropolitan Region (for place of residence), services sector (for economic sector) and public sector (for public or private sector).
Standard errors are shown in brackets. *** $p<0.001$; ** $p<0.01$; * $p<0.05$, † $p<0.1$.

In general, the data shown in the above tables indicate that class and gender have a sizeable impact on objective job quality and on people's perceptions of job quality. With a single exception (the effect of gender on perceived control over work processes), the results are in keeping with hypothesis 1. It is noteworthy that the R-squared values for the motivation models are lower than they are for the other models (the most complete model has an R-squared of only 8%). This suggests that, by comparison to the other job quality indicators, motivation is influenced by more variables that have not been included in the regressions.

Although these data indicate how gender, on the one hand, and class, on the other, generate substantial differences in job quality, they do not show how gender and class intersect in terms of the generation of inequality. In order to explore this effect and compare our results with hypothesis 2, we decided to run the regressions for men and women separately.

2. The impact of gender on the structure of class-based inequality in the labour market

Table 5 gives the coefficients for the determinants of objective job quality, disaggregated by the survey respondents' gender. An examination of the coefficients for each class-based category shows up notable differences. For example, in model 2, there are class categories, such as "unskilled workers," that generate a significant difference for men but not for women. In addition, class categories such as "skilled workers" yield a negative sign for men but a positive one for women (unlike the result for men, women's membership in that class is associated with a higher score on the objective job-quality index).

The situation is somewhat similar for "unskilled supervisors" and "works in the private sector;" in these categories, the coefficients for men are not significant or have a negative sign, but the coefficients for women are positive and significant. This does not mean that skilled women workers, women supervisors or women who work in the private sector have better-quality jobs than men, but rather that the effect associated with being a skilled worker, a supervisor or a person who works in the private sector is positive only in the case of women.

There could be a number of different reasons for this. With respect to the coefficients for social class, the data could be showing how gender gives rise to differences among the members of a given class, and especially in certain specific social classes, such as the reference category (small-scale employers). When the average scores on the objective job quality index for each gender within each class-based category are compared, it turns out that men in all social classes score higher than women (e.g. while the average score for skilled male workers was 58, it was 51 for skilled female workers). Even more importantly, the data also show that the biggest difference between men's and women's scores is for the category of small-scale employers, with men averaging a score of 71 points versus 52 points for women.

This explains why the performance of a skilled job has a positive effect for women: unlike men, for women there is a fairly sizeable improvement in job quality relative to the level of job quality recorded for the reference category when they become an unskilled supervisor or a skilled worker. This also explains why being an unskilled worker does not translate, for women, into a sizeable reduction in job quality, since, in this case, the data show that the quality of employment for small-scale female employers is quite similar to what it is for low-skilled wage earners.

In the case of employment in the private sector, something quite similar may be occurring. According to the data, whereas, in the case of men, job quality in the private sector is lower than it is in the public sector (men in the former sector have a score of 53, on average, versus one of 70 for men in the latter), the situation is the opposite for women. The average score on the job quality index for women who work in the private sector is 46, while it is 43 for women working in the public sector. Although the differential is smaller in the case of women, it is nonetheless statistically significant (The analysis of variance (ANOVA) yields a result of $p < 0.05$). Accordingly, it is no surprise that the effect of working in the private sector is positive only for women.

Table 5
Coefficients of determinants of subjective job quality in Chile, by gender

	Men		Women	
	Model 1	Model 2	Model 1	Model 2
Social class (reference: small-scale employers)				
2. Formal petty bourgeoisie	-23.46***	-22.20***	-12.27**	-9.314*
	(2.761)	(2.622)	(3.931)	(3.739)
3. Informal petty bourgeoisie	-31.65***	-23.58***	-24.07***	-15.11***
	(1.975)	(1.928)	(3.134)	(3.027)
4. Expert managers	6.114**	6.310**	15.43***	14.01***
	(2.115)	(2.030)	(3.291)	(3.144)
5. Non-managerial experts	5.959*	0.480	21.85***	11.97***
	(2.911)	(2.857)	(3.395)	(3.389)
6. Skilled workers	-12.30***	-5.884**	-1.525	5.960*
	(2.020)	(1.961)	(3.130)	(3.010)
7. Unskilled managers	-4.947	-0.229	-6.818	2.600
	(3.861)	(3.673)	(6.745)	(6.428)
8. Unskilled supervisors	-3.463	0.601	6.469†	13.89***
	(2.181)	(2.111)	(3.408)	(3.273)
9. Unskilled workers	-17.00***	-9.422***	-9.206**	1.866
	(1.932)	(1.894)	(3.060)	(2.975)
Demographic variables				
Age		0.132***		0.143***
		(0.0201)		(0.0291)
Secondary education		9.790***		8.257***
		(0.616)		(0.865)
Vocational institute/technical training centre		13.75***		13.92***
		(1.050)		(1.312)
University or higher		19.68***		27.56***
		(1.473)		(1.850)
Residence in the Metropolitan Region		1.820**		2.067*
		(0.681)		(0.828)
Economic variables				
Agricultural sector		-5.889***		-5.690***
		(0.719)		(1.237)
Manufacturing sector		2.040**		0.859
		(0.600)		(1.257)
Private sector		-2.742*		6.128***
		(1.238)		(0.899)
Constant	70.09***	53.76***	52.27***	25.39***
	(1.890)	(2.556)	(3.016)	(3.523)
Adjusted R^2	0.23	0.31	0.26	0.33
N	5 591	5 573	3 303	3 295

Source: Prepared by the authors, on the basis of data from the National Survey on Employment, Work, Health and Quality of Life of Workers in Chile (ENETS 2009-2010).
Note: Unstandardized coefficients, ordinary least squares regression.
The omitted variables are male (for gender), employers (for social class), primary education (for level of education), does not live in the Metropolitan Region (for place of residence), services sector (for economic sector) and public sector (for public or private sector).
Standard errors are shown in brackets. *** p<0.001; ** p<0.01; * p<0.05; † p<0.1.

Table 6 gives the determinants for the two subjective job-quality indices, disaggregated by gender. The models for on-the-job motivation indicate that, while the impact of class is significant for men and women (in both cases, membership in a subordinate class, such as the working class, is associated with lower levels of motivation), it tends to be weaker for women. An analysis of model 2 shows, for example, that, for women, only the coefficient for the working class is significant, and the coefficient is slightly lower than it is for men.

In other words, while for men the results are more staggered, as was to be expected in line with the hypotheses used here, in the case of women, levels of motivation tend to be similar to those seen in the reference category, except among working-class women. The survey data indicate not only that women are less motivated on the job —the average score for men is 86, while for women it is 82 (ANOVA: $p<0.000$)— but also that the average scores for each social class do not differ significantly from the score for small-scale employers (except, of course, in the case of working-class women).

The importance of gender is also reflected in the coefficients for the economic variables. For example, while employment in the manufacturing sector generates significant increases in levels of motivation for men, no such increase is seen in the case of women. In fact, on-the-job motivation among women working in the manufacturing sector does not differ significantly from the level of motivation seen in the reference category (women workers in the services sector). Similarly, employment in the agricultural sector is significant only for women, with women who work in that sector being significantly less motivated than women who work in the services sector.

The data for the agricultural sector probably reflect the fact that substandard jobs in that sector are disproportionately held by women (e.g. female seasonal workers) (Caro, 2012). By the same token, it is likely that the results for employment in the manufacturing sector for men is reflecting the fact that declining job quality and stability in the services sector is having a greater influence on how job quality is perceived by men than by women, which does not mean that women have better-quality jobs than their male counterparts. Instead, this could well be an outcome of the fact that many men —especially those in the working class— have directly experienced the transition from a manufacturing-based economy (associated with more job stability) to a service-based one (associated with less job stability and greater flexibility). In other words, unlike the large number of women who entered the formal labour market during the 1980s, when the services sector was burgeoning, it is possible that many men who now work in services are comparing the jobs that they held during the import-substitution phase (when many were employed in the manufacturing sector) with the jobs that they hold now, which are a direct result of the introduction of the neoliberal economic regime. This point of comparison could well be the underlying cause of the gender-based variations observed in the economic-sector effect on motivation on the job.

The data reveal similar trends in terms of the distinction between employment in the public and private sectors. They indicate that men employed in the private sector are significantly less motivated than their public-sector counterparts. This does not occur in the case of women, however. As was seen in the case of employment in the manufacturing sector, it is probable that these data are a reflection of the transition from an economy that provided more protected forms of employment to certain segments of workers (public-sector employees, for example) to an economy in which the labour market is more flexible. As noted earlier, this transition has probably had a greater influence on men's career paths than on those of women, most of whom joined the labour market later, during a time when more flexible employment regimes were already in place.

Table 6 also shows the results for determinants of the perception of control over work processes. As in the preceding cases, the data reflect considerable differences between men and women. First, although the general patterns for the social class variable are similar for the two, the coefficients for some categories (e.g. the working class) are much higher for men. For example, while the impact

of belonging to the working class translates into a reduction of nearly 25 points on the index for the perception of control over working processes for women, the impact of working-class membership is -38 points for men (model 2).

Second, an examination of the R-squared values for model 1 clearly shows that the explanatory power of social class is greater for men than it is for women. One possible explanation for this is that, if social class has been construed as a "masculine" attribute (Acker, 2006), then it could have a stronger impact on men's perceptions than those of women.

Although certainly worthy of consideration, this interpretation should be taken with a grain of salt because, if social class were a primarily viewed as a masculine attribute, then its impact as expressed in R-squared values should be greater for men in all of the dimensions of job quality analysed here. However, this is not the case. Another possible explanation may be that it is not social class per se, but rather the positions of authority within the production process that are "masculinized." The results of the National Survey on Employment, Work, Health and Quality of Life of Workers in Chile provide support for this conclusion, since they indicate that 37% of men hold positions of authority (manager, supervisor or any other position in which the incumbent wields control over other workers), but that only 28% of women do so. These gender-based distinctions could be having an effect whereby those who do not hold positions of authority but who may be more likely to attain such positions in the future (male workers) perceive the distinction between having control over work processes and not having that control more keenly.

The data also indicate that employment in the agricultural sector is associated with a greater degree of perceived control over work processes for men than for women. There may be a dual explanation for this gender-based difference. On the one hand, women in this sector may be more exposed to a more "top-down" control structure because of the types of jobs they perform. On the other, the jobs held by men in this sector may be comparatively more autonomous. This is not necessarily because their forms of employment are of better quality or more stable; they may simply involve more independent forms of production (e.g. the farming of small, individual plots of land). Tabulations of the results of the National Survey on Employment, Work, Health and Quality of Life of Workers in Chile provide some support for this hypothesis in that they indicate that 27% of men in the category of "informal petty bourgeoisie" are employed in the agricultural sector, whereas only 11% of the women in that category are employed in this sector of the economy.

Gender-based differences are also observed in relation to employment in the private sector, which has a negative impact on the perception of control over work processes for women but not for men. This may be because workers in the private sector are more exposed to gender-based forms of labour inequality than employees in the public sector are (Tokman, 2011). One factor may be that there are highly standardized and codified administrative regulations governing promotions in the public sector. These standardized procedures, which map out specific career paths, may leave less scope for discretionary decisions on the part of supervisors or managers regarding who they will promote and what employee attributes should be assessed positively. The fact that less scope exists for discretionary decisions does not necessarily mean that there are more women in positions of authority in the public sector, but simply that the gateways to promotion in that sector are less subject to implicit value decisions regarding "commitment or loyalty to the firm" than they are in the private sector. Those kinds of implicit assessments are much more closely associated with the male-defined concept of "good workers;" in other words, they are defined by a concept of employment that does not take into account the way in which women, unlike men, are faced with the trade-off between being the "ideal worker" (i.e. being totally committed to one's job) and fulfilling their expected role as mothers and/or housewives (Blair-Loy, 2003). It is probable that the greater importance placed on gender-based assessments of such attributes in the private sector is the reason why women who work in that sector have a lower perceived level of control over work processes than men in that sector do.

All of the regressions indicate that hypothesis 2 should be rejected. While gender is certainly a central factor in understanding the differences that exist among different segments of workers in the labour market, the role that it appears to play does not dovetail with the propositions of that hypothesis. All results for all three of the indices indicate that gender does not amplify the social-class effect, which is negative, for example, in the case of the observed quality of jobs held by people in subordinate classes.

Thus, for example, the gender effect as measured by the objective job-quality index was just the opposite of what was expected. In the "skilled workers" and "unskilled workers" categories, the effect of social class was positive but not significant for women, whereas —according to the operative hypothesis— it was expected to be statistically significant, negative and greater than in the case of men. In contrast, in many cases gender did have an amplification effect, in a positive sense, in higher-ranking classes, such as that of "expert managers." In fact, contrary to expectations, the positive effect of belonging to the class of "expert managers" was greater for women than for men.

The coefficients for men and women on the second index (on-the-job motivation) were generally similar, although the class effect was weaker for women. In this case, although the only class-based category in which the results were significant for women (the working class) yielded coefficients that were as expected, the impact was quantitatively smaller than it was in the case of men.

Finally, the results on the third index (perception of control over work processes) indicate that, although the coefficients are similar for men and women, the effect of belonging to a subordinate class was stronger for men than for women. The models thus show that it is working-class men, rather than working-class women, who have the lowest perceptions of control over work processes.

Table 6
Coefficients of determinants of subjective job quality in Chile, by gender

| | Motivation on the job | | | | Perception of control over work processes | | | |
| | Men | | Women | | Men | | Women | |
	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2
Social class (reference: small-scale employers)								
2. Formal petty bourgeoisie	-4.843†	-4.359†	2.962	3.411	-1.962	-1.477	7.440	8.115
	(2.636)	(2.607)	(4.219)	(4.170)	(3.960)	(3.959)	(5.758)	(5.631)
3. Informal petty bourgeoisie	-6.365**	-6.615**	-5.112	-3.498	-0.677	-0.420	8.491†	10.86*
	(1.880)	(1.916)	(3.308)	(3.317)	(2.822)	(2.903)	(4.566)	(4.527)
4. Expert managers	-2.867	-3.382†	1.526	1.868	-17.96***	-17.13***	-12.16*	-14.63**
	(2.015)	(2.010)	(3.470)	(3.448)	(3.024)	(3.053)	(4.781)	(4.700)
5. Non-managerial experts	-5.047†	-4.943†	2.650	0.861	-27.21***	-27.56***	-22.69***	-28.32***
	(2.795)	(2.847)	(3.592)	(3.738)	(4.195)	(4.320)	(4.942)	(5.089)
6. Skilled workers	-11.85***	-11.03***	-3.786	-1.276	-36.79***	-35.36***	-28.58***	-28.11***
	(1.923)	(1.949)	(3.298)	(3.300)	(2.882)	(2.950)	(4.554)	(4.503)
7. Unskilled managers	-3.775	-3.543	-5.904	-3.163	-0.417	0.913	8.750	10.07
	(3.588)	(3.563)	(7.570)	(7.495)	(5.360)	(5.372)	(10.28)	(10.07)
8. Unskilled supervisors	-8.785***	-8.757***	-4.886	-1.364	-24.95***	-23.11***	-12.67*	-9.952*
	(2.085)	(2.106)	(3.625)	(3.620)	(3.122)	(3.183)	(4.990)	(4.926)
9. Unskilled workers	-14.28***	-14.03***	-13.28***	-9.779**	-39.43***	-37.84***	-25.56***	-25.01***
	(1.836)	(1.880)	(3.220)	(3.256)	(2.758)	(2.851)	(4.449)	(4.445)

Table 6 (concluded)

| | Motivation on the job | | | | Perception of control over work processes | | | |
| | Men | | Women | | Men | | Women | |
	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2	Model 1	Model 2
Demographic variables								
Age		0.127***		0.235***		0.121***		0.186***
		(0.0205)		(0.0329)		(0.0304)		(0.0442)
Secondary education		1.135†		0.140		0.829		-0.747
		(0.628)		(0.984)		(0.935)		(1.322)
Vocational institute/ technical training centre		0.430		3.351*		2.087		2.833
		(1.064)		(1.472)		(1.586)		(1.976)
University or higher		2.277		4.364*		3.314		5.274†
		(1.480)		(2.067)		(2.221)		(2.778)
Residence in the Metropolitan Region		-6.054***		-4.544***		-4.104***		-4.828***
		(0.672)		(0.922)		(1.001)		(1.237)
Economic variables								
Agricultural sector		0.764		-4.284**		2.495*		-6.970***
		(0.733)		(1.402)		(1.093)		(1.875)
Manufacturing sector		3.295***		-0.647		-1.331		-0.994
		(0.608)		(1.414)		(0.906)		(1.897)
Private sector		-2.869**		-0.199		0.314		-11.07***
		(1.253)		(1.013)		(1.871)		(1.359)
Constant	96.51***	92.61***	90.07***	78.55***	83.15***	76.30***	72.50***	74.32***
	(1.794)	(2.570)	(3.171)	(3.894)	(2.698)	(3.870)	(4.383)	(5.290)
Adjusted R^2	0.05	0.07	0.06	0.09	0.25	0.25	0.17	0.21
N	5 606	5 588	3 280	3 272	5 557	5 538	3 265	3 257

Source: Prepared by the authors, on the basis of data from the National Survey on Employment, Work, Health and Quality of Life of Workers in Chile (ENETS 2009-2010).
Note: Unstandardized coefficients, ordinary least squares regression.
The omitted variables are male (for gender), employers (for social class), primary education (for level of education), does not live in the Metropolitan Region (for place of residence), services sector (for economic sector) and public sector (for public or private sector).
Standard errors are shown in brackets. *** $p<0.001$; ** $p<0.01$; * $p<0.05$; † $p<0.1$.

VII. Conclusions

The findings of this study demonstrate that class and gender have a significant impact on job quality and on people's subjective perceptions of job quality in the Chilean labour market. They also show that class, on the one hand, and gender, on the other, give rise to substantial variations in objective job quality and in on-the-job motivation. In addition, the degree of perceived control over work processes is heavily influenced by social class, but not by gender. Accordingly, evidence has been found that corroborates the general hypothesis on which this study was based (hypothesis 1), except in the case of the impact of gender on the perceived level of control over work processes.

The data also indicate, however, that the second hypothesis, relating to the intersectionality of class and gender, should be ruled out, since, in the cases of two of the three dependent variables that were analysed (objective job quality and motivation on the job), the condition of being a woman

did not amplify the impact associated with being, for example, a member of the working class. And when gender did influence the impact that social class has on the dependent variable (the perception of control over work processes), it did so in a way that ran counter to the expected effect (e.g. the negative effect on the dependent variable associated with membership in the working class is greater for men than it is for women).

The foregoing does not mean that gender is not a mechanism of fundamental importance in understanding how inequality in the labour market is generated or functions. The analyses of the sample when it was divided up into separate subsamples of men and women show that the impacts of a series of economic variables vary substantially depending on whether a worker is male or female, and this is also demonstrated by the data on the differences in working conditions and conditions of employment (Directorate of Labour, 2012).

Overall, these results suggest that both class and gender play a central role in the perpetuation of inequalities in the Chilean labour market. They also indicate that, in some cases, gender is a fundamental consideration in understanding how class-based inequalities operate or, in this specific case, how social class influences subjective aspects of job quality, such as the perception of control over work processes. In other words, in line with various gender studies dealing with industrialized nations (Ferree and Hall, 1996; McCall, 2001; Acker, 2006; Mintz and Krymkowski, 2010), these data suggest that, at a concrete level of analysis, studies on class-based inequalities should include gender as a variable because it is a major determinant of the way in which social class operates in capitalist societies in general (Wright, 1992) and, according to the results of this study, in Chile in particular.

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Brazil, 1981-2013: the effects of economic growth and income inequality on poverty

Francisco José Silva Tabosa, Pablo Urano de Carvalho Castelar and Guilherme Irffi

Abstract

This study analyses the impact of economic growth and income inequality on poverty in Brazil in the years from 1981 to 2013. A dynamic panel model was used, estimated by the two-step generalized method-of-moments system developed by Blundell-Bond (1998), in order to analyse three scenarios: the first corresponds to the entire period covered by this study (i.e. 1981-2013); the second encompasses the years from 1981 to 1994 (the period leading up to the Real Plan); and the third is the period from 1995 to 2013 (the years following the implementation of the Real Plan). The results indicate that economic growth policies that promote an increase in income in conjunction with a reduction in income disparities are more effective in combating poverty in Brazil than those that focus only on raising mean income levels. The findings also point to the existence of a pro-poor form of growth in the period following the Real Plan.

Keywords

JEL classification

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I.   Introduction

Poverty reduction is closely linked to income inequality and mean income levels in a given country or region (Bourguignon, 2002). It is therefore important to determine what type of policy will be the most effective in reversing poverty as rapidly as possible. The question, then, is: Should poverty-reduction policies focus on raising mean income levels or on decreasing income inequality?

In an effort to answer that question, a number of studies have been conducted in recent years on the effects of changes in income levels and in inequality on poverty rates, since economic growth alone has been shown to be ineffective in combating poverty in various countries (and regions) (Ravallion, 1997; Rocha, 2006).

Given that changes in poverty rates may be attributable to a redistribution of income or to economic growth (or both), it is important to gauge the effect that changes in each of these variables have on poverty levels (Ravallion, 1997). Based on the results of this study, an effort is made to identify the most influential factors in terms of poverty reduction.

Ravallion and Chen (1997), for example, estimated the income elasticity of poverty (measured on the basis of the number of persons with incomes of less than US$ 1 per day) in a sample of developing countries and obtained a value of -3. This means that, for each 1% increase in the mean income level, the percentage of persons with incomes below the poverty line shrinks by 3%.

However, there is as yet no consensus about the nature of the interrelationships between poverty, growth and income inequality. According to Barreto (2005), much of the empirical evidence appears to indicate that low-income countries in which there is a low level of inequality will respond better to economic growth policies, whereas those policies will be less effective in higher-income economies with greater levels of inequality; in the latter, policies aimed at reducing inequality will be more effective.

A series of studies on the situation in Brazil[1] show that poverty levels are more sensitive to policies designed to reduce income concentration than they are to policies aimed at boosting mean income levels. Determining how each of these factors affects the others will therefore make a highly significant contribution to the debate about what types of public policies are the most effective in reducing poverty and inequality.

The study done by Barros and others (2007) confirms that the poverty rate in Brazil is higher than it is in most countries with similar per capita income levels and demonstrates that income inequality is responsible for the fact that economic growth is relatively ineffective in reducing poverty. In other words, the poverty-reducing effect of economic growth is weaker in Brazil than it is in other countries with similar income levels.

According to Rodrik (2000), most of the policies adopted for this purpose focus on spurring economic growth, as an increase in an economy's mean income level and/or a reduction in income inequality, or a combination of the two, can lower poverty rates. It is therefore essential to gauge how much weight should be given to each of these strategies, both at the regional and state levels.

This study uses panel data models to analyse the effects of economic growth and income inequality on poverty rates in Brazil in the period from 1981 to 2013. The persistence of poverty dynamics in Brazil can be examined and estimated by drawing on pooled data and applying ordinary least squares (OLS) based on the within-group method and the generalized method of moments (system GMM), a dynamic panel model estimated using the two-step generalized method of moments or two-step GMM developed by Blundell-Bond (1998).

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[1]   See, for example, Hoffmann (2005), De Lima, Barreto and Marinho (2003) and Menezes and Pinto (2005).

The main contribution made by this study consists of estimations of model parameters for the subperiods 1981-1994 and 1995-2013 for use in analysing the situation in the periods leading up to and following the Real Plan. These findings can then be used to compare inequality and income elasticities of poverty and extreme poverty in different periods of the economic history of Brazil.

In order to achieve these objectives, this study is divided into five sections, one of which is this Introduction. In the second section, a brief overview is presented of the literature on the impact that economic growth, on the one hand, and income inequality, on the other, have on poverty. This is followed by a description of the database, models and econometric methodology that have been used. An analysis and discussion of the results of the model estimations are presented in the fourth section. The fifth and last section offers concluding remarks.

II. Literature review: the poverty, economic growth and inequality triangle

According to Bourguignon (2002), a triangular relationship exists between poverty, economic growth and income inequality whereby their interaction provides the necessary conditions for determining how great an impact an increase in mean incomes or a reduction in inequality will have in terms of poverty reduction. For this study, it seemed fitting to present the empirical evidence gathered in three sections.

The first is devoted to studies on the relationship between poverty and economic growth. The second focuses on an analysis of the relationship between poverty and inequality. The third covers the literature on the relationship between economic growth and income inequality.

1. Poverty and economic growth

In the international and national empirical literature on the relationship between economic growth and poverty, two different measurements are used to gauge economic growth: gross domestic product (GDP) and mean income. These studies' findings corroborate each other to the extent that they all indicate that, as noted by Araújo, Tabosa and Khan (2012), there are two fundamental factors at work: the average growth rate and the initial level of income inequality.

The relationship between economic growth and poverty reduction can be measured by estimating the corresponding income elasticity or growth elasticity. If it is high, then public policies designed to combat poverty based on economic growth will be more effective since, in this type of situation, boosting income will have a proportionally greater effect in terms of poverty reduction. On the other hand, if the coefficient of elasticity is low, the most appropriate poverty reduction strategies will be those that combine economic growth with some type of income redistribution (Marinho and Araújo, 2010).

Ravallion and Chen (1997) have estimated income-poverty and income-inequality elasticities on the basis of data from 45 countries. Their results indicate that, if the income level were to rise by 1% in countries with low levels of inequality, the poverty rate would fall by 4.3%, whereas, in countries with high levels of inequality, the same increase in mean income would lead to a 0.6% drop in the poverty rate. They thus conclude that, in this latter case, growth has no more than a weak impact on poverty levels. However, if the extent of inequality declines as a result of economic growth, then a greater reduction in poverty will be achieved.

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Chen and Wang (2001) studied the relationships between poverty, income and inequality in China in the 1990s and reached the conclusion that economic growth helped to lower poverty levels but that income concentration helped to raise them.

More recently, Figueiredo and Laurini (2015) undertook an analysis of the elasticity of poverty on an international scale. They seek to apply a non-parametric method to a country panel using the World Bank's PovcalNet database, which includes 139 observations for 93 developing countries. The authors' chief concern was to correct for any problems of endogeneity that might arise as a result of a series of factors, such as simultaneity, financial forces that have the same kind of effect on the components being analysed or the hypothetical situation that is usually assumed to exist in terms of the functional relationships between growth, poverty and inequality (a hypothesis according to which the relationship is linear, whereas there are some indications in the literature that this is not the case).

These authors contend that another contribution made by their work is a determination of the direct and indirect effects of growth on poverty and on inequality based on the non-parametric methodology that they have employed. Their findings indicate that the methodology that is generally used in empirical studies on growth, poverty and inequality usually overestimates the effects of growth on poverty. This implies that economic crises or economic booms have less of an influence on poverty than they are generally assumed to have.

At the national level, using data for 26 Brazilian states covering the period from 1985 to 1999 and a methodological procedure that makes it possible to disaggregate the variation in poverty levels deriving from changes in mean income levels and modifications in income concentration (measured by the Gini coefficient), Marinho and Soares (2003) find proof of the effectiveness of growth as a tool for poverty reduction, especially in the northern states of Brazil. This supports the focus on raising incomes which has been the preferred strategy for fighting poverty in Brazil.

In another work on this topic for Brazil, Hoffmann (2005) uses the database for the 1999 National Household Survey (PNAD) and finds that a 1% increase in per capita household income in Brazil leads to a 0.84% reduction in poverty and that the absolute value of the elasticity of poverty rises in step with income and decreases in step with increases in inequality.

Manso, Barreto and Tebaldi (2006) used the PNAD data for 1995-2004 as a source of evidence on the relationships between rising income levels, reductions in poverty levels and the wealth distribution profile. They find that the growth of mean income and income distribution are sufficient to account for a large part of the variations in poverty rates observed at the state level in Brazil. The results of these studies thus demonstrate that growth-oriented tools for combating poverty are more effective when they are combined with income redistribution (Barreto, 2005).

Another point made in a number of studies on the relationship between poverty and economic growth is the importance of determining whether growth is pro-poor or not. Ravallion (2004) distinguishes between two types of pro-poor growth: (i) growth that entails a greater reduction in poverty than would be seen if everyone's incomes had risen by the same proportion (Kakwani and Pernia, 2000), and (ii) growth that reduces poverty, regardless of the proportions involved (Ravallion and Chen, 2003).

In the first case, an analysis can be conducted on the basis of the income elasticity of poverty; if that elasticity is greater than 1 (in absolute terms), then the reduction in poverty will outpace the growth of income (Kakwani and Pernia, 2000). In the second case, regardless of the value obtained for the income elasticity of poverty, if the increase in income translates into a reduction in poverty, then growth is pro-poor (Ravallion, 2004).

The study carried out by França (2010) on pro-poor growth and its regional effects in Brazil is a specific example of this type of analysis and has an objective that is quite similar to the aim of this paper. The results that he obtained using the methodology developed by Kakwani, Khandker

and Son (2004) to analyse data on the various regions of Brazil for the period 1995-2005 indicate that anti-poverty policies have had a greater impact in the southern and south-eastern parts of the country. The author then calculated the income elasticity and the inequality elasticity of poverty and found that, since the latter was greater than the former, the reduction in inequality as measured by the Gini coefficient was more sensitive to reductions in the level of inequality than to increases in mean income. He then used the methodology developed by Kraay (2004) to determine how much influence is exerted by each component of pro-poor growth and found that increases in income have been relatively more influential than changes in inequality as a source of pro-poor growth in the north-eastern region of the country.

Kakwani, Neri and Son (2010) used PNAD data for 1995-2004 and a poverty line that takes into account the cost of living in the various regions of the country[2] to analyse trends in social indicators based on mean per capita income levels. Their findings indicated that growth in Brazil in the post-Real Plan period was indeed pro-poor.

Netto Júnior and Figueiredo (2014) used a non-parametric approach to analyse pro-poor growth in Brazil and in its main geographic regions and states in the periods 1987-1993, 1993-1999 and 1999-2007. Their results suggest that all parts of the country experienced pro-poor growth in 1987-2007.

2. Poverty and inequality

Income inequality is an important aspect of the poverty debate. Poverty is a worldwide problem, and in some countries it persists despite the presence of growing stocks of material wealth. The scope and severity of this problem are demonstrated by the number of persons who are living in poverty in all of the world's countries.

An intrinsic link exists between inequality and poverty. According to Bourguignon (2002) and Ravallion (1997), a decrease in income inequality is an important poverty-reduction tool, whereas economic growth may not be so crucial. They thus conclude that reductions in inequality can play a critical role in ensuring that growth will actually lead to a decline in poverty levels.

Banerjee and Duflo (2003) draw attention to the fact that the relationship between economic growth and inequality is not linear since inequality has both direct and indirect effects on poverty. Figueiredo and Laurini (2015) sought to measure those effects using a panel of 93 countries. These authors confirm the existence of a non-linear association between growth and inequality. They argue that not enough attention is devoted to that relationship in most studies on the subject and that it has a positive link to poverty.

Along these same lines, it has been shown that poverty can be rapidly reduced when a country whose economy is growing has a less unequal distribution of income (Barreto, 2005). Accordingly, in addition to attaining their specific goal, public policies aimed at reducing inequality can also indirectly achieve other economic policy objectives, such as more rapid growth and reductions in poverty. Since poverty is more sensitive to inequality than it is to growth (Marinho and Soares, 2003), anti-poverty policies should focus on methods for diminishing income concentration.

According to Rocha (2006), the poverty rate in Brazil was cut by around two percentage points in 2001-2004 thanks to a number of factors (e.g. changes in the distribution of labour income and increases in social benefits) whose effects varied across regions. This author believes that the persistence of poverty in Brazil is largely attributable to the existing level of inequality. Rocha contends that poverty can be cut both by boosting income and by improving income distribution and that the

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[2]   For further details, see Kakwani, Neri and Son (2010) and Ferreira, Lanjouw and Neri (2003).

emphasis should be placed on reducing income inequality, because raising income levels without reducing inequality simply defers the challenge of putting an end to poverty to some point of time in the future.

A very high initial level of inequality may dampen the poverty-reduction effects of economic growth, although, based on a review of studies on growth and inequality, Ferreira (1999) concludes that no consensus yet exists as to whether or not growth is influenced by the initial level of income inequality.

In the past few decades, poverty has either declined very little or actually risen in a considerable number of countries. For Fosu (2010), this disappointing outcome is partially attributable to the combination of a concentration of growth and rising income inequality in many Latin American countries.

3. Growth and inequality

The literature includes analyses of the relationship between economic growth and inequality that take the causalities existing between these variables into consideration. A wide range of related issues have become subjects of debate, including the causes of inequality, the mechanisms by which inequality is perpetuated over time and the relationship between inequality and the economic development process. Diniz (2005) posits the existence of dual causality between these two variables.

Kuznets' (1955) inverted-U hypothesis serves as the starting point for this line of reasoning, according to which, at first, inequality will increase as economic development takes place, which occurs when an economy makes the transition from a rural model to industrialization (transferring labour from the least productive sector to the most productive one). Later on, inequality will decline once the majority of the workforce is employed in the most productive sector. According to this thesis, development policy could focus entirely on promoting economic growth, since such growth would ultimately lead to a decrease in inequality. In a comparative study of 98 countries, Barro (2000) shows that the Kuznets curve exhibits a certain degree of empirical regularity over time, but that it does relatively little to explain the variation in inequality across countries.

Some authors believe that inequality may be prejudicial to economic growth. As summarized by Castelar (2007), empirical studies, especially those conducted since 1996, show that high initial levels of inequality curb subsequent economic growth. Most of the literature on the subject indicates that a one standard deviation decrease in inequality boosts the annual per capita GDP growth rate by between 0.5 and 0.8 percentage points. However, when using panel data, the relationship between income inequality and economic growth appears to be weaker, which suggests that the empirical regularity seen in the Kuznets curve is robust for cross-sectional data but disappears when fixed country effects are introduced. Examples of these studies include those of Fields and Jakubson (1994), Fishlow (1972), Deineger and Squire (1998) and Barreto, Melo Neto and Tebaldi (2001).

According to Stewart (2000) and Fosu (2010), a high level of income inequality gives rise to political instability, uncertainty, less investment, slower growth and populist redistributive tax policies. Deterrence effects and greater inequality influence the richer strata of society, which press for preferential tax treatment. This, in turn, leads to over-investment in certain areas and curbs growth.

Easterly (2000) and Adams (2004) analysed the relationships between growth and income inequality across countries and observed that the growth elasticity of poverty is greater in countries with lower Gini coefficients (less inequality). These authors underscore the importance of a reduction in inequality in determining the response capacity of poverty to rising income levels.

On the other hand, studies such as those of Li and Zou (1998) and Forbes (2000), which used five-year means for 35-country samples, indicate that, when panel data are employed, the negative

relationship between growth and inequality disappears. Barro (2000) suggests that the negative impact of inequality on growth depends on how wealthy a country is, although this relationship is not robust.

Castelar (2007) analysed the relationship between growth and inequality in the states of Brazil on the basis of panel data for 1985-2002. Using the two-step Arellano and Bond estimator, he found that income inequality depresses economic growth and was able to corroborate the convergence hypothesis, according to which initially low income or output levels are correlated with higher growth rates.

In an analysis of Brazil, Ferreira and Cruz (2010) use a threshold-effect model to study the existence of income-inequality convergence clubs at the municipal level in 1991-2000. These authors detected six such clubs, with the factors that reduced the inequality of income distribution in the municipalities of Brazil having asymmetric effects. In the course of this convergence process, labour income was more influential than income derived from government transfers in the reduction of inequality.

III. Data source and description

The database used here includes information on 25 federation units (i.e. 24 states and the federal district) for 1981-2013.[3] All the data have been taken from the National Household Survey (PNAD) conducted by the Brazilian Geographical and Statistical Institute (IBGE).

The income variable is per capita household income (*RM*) and is estimated on the basis of the ratio between total household income and the number of members of the household. We then calculate the arithmetic mean of that variable in order to obtain the mean income for each state and the federal district.[4]

Families whose per capita household incomes fall short of the sums needed to meet their basic needs are classified as poor. As an absolute poverty indicator, the metric proposed by Foster, Greer and Thorbecke (1984), known as the poverty headcount index (P_0), is used. This index is defined as follows:

$$P(\alpha) = \int_0^{PL} \left(\frac{PL - y}{PL} \right)^{\alpha} f(y) dy$$

(1)

For the poverty headcount index (P_o), $\alpha = 0$, the poverty line (*PL*) employed by the Institute of Applied Economic Research (IPEA)[5] is used, while for the extreme poverty headcount index, $\alpha = ext$, the extreme poverty line (P_{ext}) is used.[6]

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[3] Since the period of analysis stretches from 1981 to 2013, the data for the State of Tocantins, which was established in 1990, were combined with those for the State of Goiás. Thus, the calculations for income and for poverty and inequality indices for the years from 1990 on were performed by grouping the data for those two states. In addition, since the National Household Survey was not conducted in 1994 or in 2000, an interpolation of the arithmetic means for the preceding and following years was used.

[4] All of the monetary variables have been updated to reflect real 2013 values using the national consumer price index for October 2013 as a deflator.

[5] The poverty line being used here is two times the extreme poverty line.

[6] The extreme poverty line being used here is an estimate of the value of a food basket representing the minimum number of calories needed to sustain a person according to the recommended levels issued by the United Nations Food and Agriculture Organization (FAO) and the World Health Organization (WHO). The series is calculated on the basis of responses to the National Household Survey (PNAD/IBGE). For further details, see www.ipeadata.gov.br.

The Gini coefficient (G) is used as the metric for income inequality.[7] In order to calculate this coefficient, per capita household incomes have to be placed in ascending order to obtain the Lorenz curve, which plots the cumulative percentage of the population against the cumulative percentage of income for each percentile. The Gini coefficient for each state and for the federal district can then be calculated from these data.[8]

Table 1 shows that the mean poverty headcount index (P_0) was 0.3992 for the period under analysis, while the lowest and highest values were 0.0167 and 0.8616, respectively. In other words, 39.92% of the Brazilian population was below the poverty line used in this study, while 17.76% of these people were living in extreme poverty.

Mean per capita household income in 1981-2013 was 571.76 reais. The lowest amount to be registered, 160.13 reais, was for Piauí in 1983, while the highest, 1,266.63 reais was for São Paulo in 2012. The mean value for inequality (expressed in terms of the Gini coefficient) was 0.5585, with the range running from 0.3933 (Roraima in 1983) to 0.6665 (Piauí in 1990).

Table 1
Brazil: descriptive statistics for all states, 1981-2013

Variables	Observations	Mean	Standard deviation	Lowest	Highest
Poverty headcount index (P_0)	800	0.3992	0.1885	0.0167	0.8616
Extreme poverty headcount index (P_{ext})	800	0.1776	0.1309	0.0071	0.6279
Per capita household income (Rm)	800	571.76	225.56	160.13	1 266.63
Gini coefficient (G)	800	0.5585	0.0440	0.3933	0.6665

Source: Prepared by the authors.

1. The econometric model

In order to quantify the effects of economic growth and income inequality in terms of poverty, a dynamic panel data model is used:

$$ln\left(P_{k,it}\right) = \beta_0 + \beta_1 \, ln\left(P_{k,it-1}\right) + \beta_2 \, ln\left(Rm_{i,t}\right) + \beta_3 \, ln\left(G_{i,t}\right) + v_t + u_{i,t} \tag{2}$$

Where $P_{k,it}$ represents the poverty index k based on the poverty line (i.e. $k = 0$ for the poverty headcount index and $k = ext$ for the extreme poverty headcount index); $Rm_{i,t}$ corresponds to mean per capita income; $G_{i,t}$ is the Gini coefficient; v_t stands for the unobservable fixed effects for individuals; and $u_{i,t}$ is the idiosyncratic error term. The i and t subindices refer to the federation unit (state or federal district) and the year.

The model is specified by logarithms that yield the income and inequality elasticities of poverty, which are represented by the coefficients β_2 and β_3, respectively.

The specification of the model (2) is based on the assumption that the existing level of poverty tends to influence poverty dynamics. This justifies the introduction of a one-period lagged independent variable $\left(P_{k,it-1}\right)$ as an explanatory variable.[9]

The hypotheses for the model are that $E\left[v_i\right] = E\left[u_{i,t}\right] = E\left[v_i u_{i,t}\right] = 0$ for $i=1,...,N$ and $t=1,...,T$; the error is not temporally correlated or, in other words, $E\left[u_{i,T} u_{i,S}\right] = 0$ for $i=1,...,N$ and

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[7] According to Litchfield (1999), this index fulfils four of the five axioms that inequality measures are generally required to meet: the Pigou-Dalton transfer principle, the income scale independence principle, the principle of population and the symmetry (or anonymity) principle. The axiom that it does not meet is the additive decomposability principle.

[8] The Gini coefficient ranges from 0 (zero) to 1 (one). The closer it is to 1, the greater the degree of income inequality, while the closer it is to zero, the less inequality there is (Hoffmann, 1998).

[9] For evidence regarding the perpetuation of poverty in Brazil, see Ribas, Machado and Golgher (2006).

$\forall t \neq s$. In addition, the initial condition of $E\left[P_{k,it}u_{i,t}\right]=0$ for i=1,...,N and t=1,...,T is imposed (Ahn and Schmidt, 1995).

An observation regarding the problem of endogeneity is called for here. Figueiredo and Laurini (2015) voice concern about the possibility of endogeneity being a factor when the usual Ravallion and Chen (1997) methodology is used to measure the effects of economic growth on poverty and inequality. According to these authors, this can arise as a result of two mechanisms: the simultaneous determination of poverty and growth, which could generate a correlation with the error component (endogeneity problems can be caused by unobserved factors that affect these two components simultaneously or by a financial development process that simultaneously affects poverty, growth and inequality); or the direct effect that a growth trend could have on the poverty measurement.

Another problem that can bias the estimates and that has also been pointed out by Figueiredo and Laurini (2015) has to do with the functional form used for the relationship between poverty, growth and inequality. With linear specifications, it is assumed that the effects of growth on poverty and inequality are constant and independent of the levels of growth and inequality. However, a number of studies on the subject have raised questions in this regard, with their authors arguing that hypothesizing a linear relationship between growth, inequality and poverty can produce a misspecification that biases the estimates because these relationships may not, in fact, be linear.

The presence of $P_{k,it-1}$ as an explanatory variable in equation (2) can give rise to endogeneity[10] and produce a bias in the dynamic panel. In that case, the OLS estimates will tend to be biased and inconsistent, overestimating the coefficient of that variable. An effort can be made to correct this bias by using the fixed effects estimator (within-groups), which generates slightly smaller standard deviations of the coefficients. However, in order to solve the problem of endogeneity, a first-differences transformation should be applied, after which an estimate can be made using the generalized method of moments (GMM) while minimizing the moment conditions of the distribution. Using this approach, equation (2) becomes:

$$\Delta ln\left(P_{k,it}\right) = \beta_1 \Delta ln\left(P_{k,it-1}\right) + \beta_2 \Delta ln\left(Rm_{i,t}\right) + \beta_3 \Delta ln\left(G_{i,t}\right) + \Delta u_{i,t} \qquad (3)$$

where Δ is an operator of differences and, given equation (3), $\Delta ln\left(P_{k,it-1}\right)$ and $\Delta u_{i,t}$ are correlated and the endogeneity persists. It thus becomes necessary to use some instrument for $\Delta ln\left(P_{k,it-1}\right)$. The hypotheses adopted in equation (2) imply that the moment conditions $E\left[\Delta ln\left(P_{k,it-s}\Delta u_{it}\right)\right]=0$ for $t=3,4,...,T$ and $s\geq2$ are valid. In line with those moments, Arellano and Bond (1991) suggest that $\Delta ln\left(P_{k,it-s}\right)$ be used for $t=3,4,...,T$ and $s\geq2$ as instruments for equation (3).

In the second case, the values of the variable that is lagged one or more periods can be regarded as valid instruments for estimating equation (3). In the last case, the values of the variables that are lagged two or more periods are valid instruments for estimating equation (2).

However, Arellano and Bover (1995) and Blundell and Bond (1998) explain that these instruments are weak when the dependent and explanatory variables exhibit a high degree of persistence, when the relative variance of the fixed effects increases or when both of these things occur. In this case, the estimator is inconsistent and biased. To solve this problem, these authors recommend estimating a system that combines the equations in differences (equation (3)) with the equations in levels (equation (2)), i.e. the system GMM.

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[10] In the event that the explanatory variables of the model are correlated with the residuals $E\left(Z_{it},\varepsilon_{it}\right)\neq 0$. Any given variable can be classified as: strictly exogenous, if it is not correlated with the past, present and future error terms; weakly exogenous, if it is correlated only with the past values of the error term; or endogenous, if it is correlated with the past, present and future error terms.

For the equations in differences, the instruments are the same as those described earlier, whereas, for the regression in levels, the appropriate instruments are the lagged differences of the corresponding variables. For example, assuming that the differences of the explanatory variables are not correlated with the individual fixed effects (for $t = 3,4,...,T$) and $E\left[\Delta ln\left(P_{k,i2}v_i\right)\right] = 0$ for i=1,2,3,...,N, the explanatory variables in differences and $\Delta ln\left(P_{k,it-1}\right)$, in the event that they are exogenous or weakly exogenous, are valid instruments for the equation in levels. The same is true if they are endogenous, but the instruments are then the explanatory variables in differences lagged one or more periods $\Delta ln\left(P_{k,it-1}\right)$.

The consistency of the system-GMM estimator depends on the absence of a serial correlation in the error term and the validity of the additional instruments. Initially, the null hypotheses of absence of autocorrelation of first and second order of the residual series need to be tested. In order for the estimators for the parameters to be consistent, the hypothesis of absence of autocorrelation of first order must be rejected and that of second order must be accepted. The Hansen and Sargan tests can then be run to check the validity of the instruments used in the system GMM.

Since the OLS estimator for the lagged dependent variable theoretically yields downward biased estimates, whereas the within-group estimators generate upward biased estimates, the appropriate estimate of the parameter of $ln\left(P_{k,it-1}\right)$ should be somewhere between the two (OLS and within group).

Following this description of the econometric and statistical procedures for arriving at these estimates, the next section presents an analysis and discussion of the results obtained using the system-GMM method to ensure that the variances of the parameters are robust to heteroscedasticity and autocorrelation. The estimator obtained by that means was corrected using the method developed by Windmeijer (2005) to avoid underestimating the true variances in the finite sample.

IV. Analysis and discussion of the results

The period 1981-2013 was chosen for the comparison of the income and inequality elasticities of poverty, along with the subperiods 1981-1994 and 1995-2013, so that the values for these elasticities in the years prior to and following the implementation of the Real Plan and in different stages in the economic development process of Brazil could be analysed.

First of all, equation (2) was estimated using OLS and the within-group method $ln(Rm_{i,t})$. In selecting the estimated model, account was taken of the results of the Hausman specification tests to check whether $Rm_{i,t}$ and $G_{i,t}$ are endogenous and the results of the Hansen and Sargan tests of the instruments' validity. These tests showed that the $Rm_{i,t}$ and $G_{i,t}$ variables did not have to be treated as endogenous variables and that the use of the second lags of the variables and the additional instruments required by the system GMM are valid.

Tests were also run to detect the presence of autocorrelation of first order of the residuals in level and in first difference (i.e. second order). The results of these tests are shown in tables 2, 3 and 4. None of the estimates rejects the null hypothesis, so the residuals exhibit a first-order correlation. In order to prevent the estimates from being biased, it is necessary to mitigate their presence. One of the ways of doing this is to estimate the model in first difference using, for example, the system-GMM estimator.

Consequently, the estimates obtained using the system-GMM method are taken into account in the analysis of the results. Tables 2, 3 and 4 show the results for the selected models that were estimated using OLS and the within-group method for the poverty headcount index (P_0) and extreme

poverty headcount index (P_{ext}) in Brazil. The OLS and within-group estimates are given only as a means of showing that the estimated value of the lagged dependent variable falls between the estimated values obtained using these two methods. This may be viewed as an indication that the bias generated by the presence of endogenous variables and unobservable fixed effects has been corrected for by using the system-GMM method.

1. 1981-2013

Table 2 shows the income and inequality elasticities for poverty and extreme poverty based on the corresponding headcount indices for the period 1981-2013. The results appear to confirm the hypothesis concerning poverty persistence, given the statistical significance of the variable $ln\left(P_{k,it-1}\right)$. In this regard, our figures corroborate the presence of poverty persistence as measured by the poverty headcount index[11] (0.1025 for P_0) and even more so in the case of the extreme poverty headcount index (0.1356 for P_{ext}). These results indicate that, while no steep upturn in poverty has been observed, poverty persistence is indeed a reality in the states of Brazil,[12] and they thus provide corroboration for the findings of Ribas, Machado and Golgher (2006), Marinho and Araújo (2010) and Marinho, Linhares and Campelo (2011).

Two of the determinants that have played a significant role in the reduction of poverty are mean per capita household income and the level of inequality as measured by the Gini coefficient.

The estimated coefficients of the income elasticity of poverty showed the expected (negative) signs, since increases in mean per capita household incomes lower poverty levels. The values of these coefficients were -0.9235 for P_0 and -1.1418 for P_{ext}. This means that a 10% increase in mean per capita household income in Brazil will translate into a 9.23% reduction in the poverty headcount index and in an 11.41% decline in the extreme poverty headcount index.

In line with the definitions used by Kakwani and Pernia (2000), the income elasticities of poverty appear to indicate that growth in Brazil has been pro-poor only in the case of extreme poverty, for which the estimated coefficient is greater than 1 (in absolute terms). In other words, the decrease in extreme poverty levels has outstripped the increase in income levels. However, in line with the definitions used by Ravallion and Chen (2003), growth has been pro-poor for both the poor and the extremely poor segments of the population, since growth has reduced poverty levels.

The estimated coefficients for the inequality elasticity of poverty also show the expected (positive) signs, as the decrease in inequality (as measured by the Gini coefficient) lowers poverty levels. The values of these coefficients were 1.9470 for P_0 and 2.8912 for P_{ext}. In other words, a 10% decrease in income inequality translates into a 19.47% reduction in the poverty headcount index and a drop of 28.91% in the extreme poverty headcount index in Brazil.

Another important finding is that the repercussions of the income effect and of the inequality effect are greater in the case of the extreme poverty headcount than they are in that of the poverty headcount. This is because, since the extreme poverty line is lower than the poverty line, it is more sensitive to upward movements in income levels and reductions in the level of inequality.

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[11] Statistically significant at 1%.

[12] According to Rocha (2006), the high poverty rate in Brazil is primarily associated with the poor distribution of resources among the country's inhabitants.

Table 2
Brazil: estimated values for the poverty headcount index and the extreme poverty
headcount index, 1981-2013

Explanatory variables	Poverty (LnP$_0$)			Extreme poverty (LnP$_{ext}$)		
	Ordinary least squares	Within-group	GMM system	Ordinary least squares	Within-group	GMM system
$lnP_{0,it-1}$	0.2001* (0.0255)	0.0862* (0.0247)	0.1025* (0.0385)	0.3596* (0.0231)	0.1254 * (0.0208)	0.1356* (0.0252)
$lnRm_{it}$	-0.6732* (0.0311)	-0.9476* (0.0325)	-0.9235* (0.1130)	-1.0271* (0.0418)	-1.3882* (0.0388)	- 1.1418* (0.0705)
lnG_{it}	1.3401* (0.1065)	2.0793* (0.1069)	1.9470* (0.2364)	2.1421 * (0.1334)	3.2094* (0.1224)	2.8912* (0.2352)
Constant	4.3633* (0.1895)	6.3059* (0.1985)	6.0746* (0.6982)	6.3827* (0.2426)	8.8043* (0.2282)	8.3485* (0.4189)
F statistic	2 262.58	1 382.37	127.22	3 156.57	2 055.37	410.18
p-value	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
R^2	0.8975	-	-	0.9306	-	-
N	775	775	775	775	775	775
Groups	25	25	25	25	25	25
H$_0$: absence of autocorrelation in residuals of first order		p-value	0.012		0.026	
H$_0$: absence of autocorrelation in residuals of second order		p-value	0.241		0.921	
Hansen test		Prob>chi2	0.450		0.367	
Sargan test		Prob>chi2	0.637		0.621	

Source: Prepared by the authors.
Note: * denotes statistical significance at 1%.

In addition, the value of the coefficient for the inequality elasticity of poverty is higher (in absolute terms) than the coefficient for the income elasticity of poverty in the case of both the poverty headcount index and the extreme poverty headcount index. This points to the possibility that the effects which reductions in inequality have in terms of decreases in poverty levels may be greater than the effects of increases in mean income levels. One possible explanation for this is that increases in income levels are transmitted to the poor segment of the population in the region in a disproportionate (or unequal) manner. These results are similar to those obtained by França (2010).

Along these same lines, Ravallion (1997 and 2004), Marinho and Soares (2003), Bourguignon (2002), López and Seven (2004), Menezes and Pinto (2005), Barreto, França and Oliveira (2008) and Marinho, Linhares and Campelo (2011) all contend that growth-oriented anti-poverty policies would be more effective if they were coupled with income redistribution measures.

As noted in the Introduction, the main contribution made by this study is its analysis of the effects of economic growth and income inequality on poverty levels in Brazil during different periods: the years leading up to the introduction of the Real Plan, when inflation was endemic in the country; and the years following its implementation, when prices stabilized. This analysis is presented in the following sections.

2.   Before the Real Plan (1981-1994)

Table 3 shows the results for the selected models that were estimated using OLS, the within-group method and the system-GMM method for the poverty headcount index (P_0) and the extreme poverty headcount index (P_{ext}) in Brazil for the period leading up to the Real Plan.

Here again, the poverty persistence hypothesis appears to be confirmed, given the statistical significance of the variable $ln\left(P_{k,it-1}\right)$, with the degree of persistence being quite high in relation to the poverty headcount index (0.1212 for P_0) and somewhat less so in the case of the extreme poverty headcount index (0.1230 for P_{ext}). These results indicate that, while poverty has not trended steeply upward in the various states that make up Brazil, it has been persistent.

As is also true for the entire period under study, the results for this subsample support the conclusion that increases in mean per capita household income and decreases in inequality as reflected in the Gini coefficient, which have been considered here as the determinants of poverty, have indeed made a significant contribution to the reduction in poverty levels.

By the same token, as mentioned in the preceding section, the coefficients for the income elasticity of poverty were negative, as expected, since increases in mean per capita household incomes reduce poverty levels. These coefficients were -0.7813 for P_0 and 0.8607 for P_{ext}. This means that a 10% rise in mean per capita household income in Brazil will translate into a 7.81% reduction in the poverty headcount index and an 8.61% drop in the extreme poverty headcount index. While for Kakwani and Pernia (2000), this does not constitute pro-poor growth, it does qualify as pro-poor growth when using the definitions advocated by Ravallion and Chen (2003), since growth has lowered poverty levels.

The estimated coefficients for the inequality elasticity of poverty are, here again, positive, as declines in inequality, as reflected in the Gini index, reduce poverty. These coefficients were 1.7487 for P_0 and 2.6541 for P_{ext}. In other words, a 10% decrease in income inequality translates into a 17.49% drop in the poverty headcount index and a 26.54% decrease in the extreme poverty headcount index in Brazil.

As before, the repercussions of the income effect and the inequality effect are greater in the case of the extreme poverty headcount index (P_{ext}) than in that of the poverty (P_0).

Table 3
Brazil: estimated values for the poverty headcount index and the extreme poverty headcount index, 1981-1994

Explanatory variables	Poverty (LnP$_0$)			Extreme poverty (LnP$_{ext}$)		
	Ordinary least squares	Within-group	GMM-system	Ordinary least squares	Within-group	GMM-system
$lnP_{0,it-1}$	0.0353* (0.0168)	0.1493* (0.0211)	0.1212* (0.0264)	0.1607* (0.0330)	0.0495*** (0.0282)	0.1230* (0.0421)
$lnRm_{it}$	-0.9131* (0.0411)	-1.1542* (0.0511)	-0.7813* (0.0705)	-1.2910* (0.0581)	-1.7918* (0.0630)	- 0.8607* (0.0791)
lnG_{it}	1.5575* (0.1483)	2.2704 * (0.1535)	1.7487* (0.3864)	2.1949* (0.1912)	3.4138 * (0.1883)	2.6541* (0.2532)
Constant	5.6512 * (0.2458)	8.6960* (0.3363)	6.2928* (0.5381)	7.6854 * (0.3349)	11.2595 * (0.4100)	8.8164* (0.4758)
F statistic	626.03	281.22	83.17	991.05	374.07	180.89
p-value	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
R^2	0.8543	-	-	0.9051	-	-
N	325	325	325	325	325	325
Groups	25	25	25	25	25	25
H$_0$: absence of autocorrelation in residuals of first order		*p*-value	0.040			0.013
H$_0$: absence of autocorrelation in residuals of second order		*p*-value	0.134			0.892
Hansen test		Prob>chi2	0.369			0.425
Sargan test		Prob>chi2	0.425			0.538

Source: Prepared by the authors.
Note: * and *** denote statistical significance at 1% and 10%, respectively.

3. After the Real Plan (1995-2013)

Table 4 shows the results for the selected models that were estimated using OLS, the within-group method and the system-GMM method for the poverty headcount index (P_0) and the extreme poverty headcount index (P_{ext}) in Brazil for the years following the Real Plan.

The poverty persistence hypothesis is confirmed in this case, too, given the statistical significance of the variable $ln\left(P_{k,it-1}\right)$. Here again, the degree of persistence is high for the poverty headcount index (0.7690 for P_0) and less so for the extreme poverty headcount index (0.5905 for P_{ext}). These results, like those obtained for the period as a whole and for the subsample for the years prior to the Real Plan, indicate that poverty in the states of Brazil has been persistent but has not risen sharply.

As in the other cases, mean per capita household income and the Gini index also made a significant contribution to reductions in poverty in the years following the Real Plan.

The estimated coefficients for the income elasticity of poverty are negative in this case as well, thereby corroborating the finding that increases in mean per capita household income reduce poverty. These coefficients were -1.1560 for P_0 and -1.3410 for P_{ext}. In other words, a 10% upturn in mean per capita household income in Brazil translates into an 11.56% reduction in the poverty headcount index and a 13.41% drop in the extreme poverty headcount index. Consequently, in this case, growth is shown to be pro-poor both for poor segments and extremely poor segments of the population, and this holds true regardless of whether the definitions espoused by Kakwani and Pernia (2000) or those proposed by Ravallion and Chen (2003) are used.

The estimated coefficients for the inequality elasticity of poverty are positive, as expected, which means that a decline in inequality (as measured by the Gini index) reduces poverty levels. These coefficients were 2.0564 for P_0 and 2.2508 for P_{ext}. In other words, a 10% decrease in income inequality translates into a 20.56% reduction in the poverty headcount index and a drop of 22.50% in the extreme poverty headcount index for Brazil.

As has also been shown to be true for the period under study as a whole and for the period 1981-1994, the repercussions of the income effect and the inequality effect are greater in relation to the extreme poverty headcount index than they are for the poverty headcount index (P_0).

The explanation for this, as noted earlier, is that, since the extreme poverty line is lower that the poverty line, it is therefore more sensitive to upward movements in income levels. In short, all the results for both subperiods (the years leading up to the Real Plan, which were marked by spiraling inflation, and the years that followed it, when inflation rates were low and stable) are similar and fit in with the results obtained for the study period as a whole. Nonetheless, the income and inequality elasticities of poverty, both for the poverty and extreme poverty headcount indices, are lower (in absolute values) for the pre-Real Plan period that for the other two periods that have been analysed.

This is probably attributable to the instability that plagued the Brazilian economy in the 1980s, when high inflation was coupled with slow growth. In this kind of situation, poverty levels were less sensitive to changes in the levels of income and inequality.

As was to be expected, the results also indicate that decreases in inequality may have a greater effect in terms of reductions in poverty rates than increases in mean income levels do. Again, one possible explanation for this is that upswings in income levels are transmitted disproportionately or unequally to the poor segments of the population. In other words, income transfer policies have not been focused on the most underprivileged members of society.

Explanatory variables	Poverty (LnPo)			Extreme poverty (LnP$_{ext}$)		
	Ordinary least squares	Within-group	GMM-system	Ordinary least squares	Within-group	GMM-system
$lnP_{0,it-1}$	0.8834* (0.0243)	0.6641* (0.0306)	0.7690* (0.0656)	0.6948* (0.0306)	0.3068* (0.0356)	0.5905* (0.0613)
$lnRm_{it}$	-0.7596* (0.0333)	-0.8641* (0.0398)	-1.1560* (0.0739)	-0.9371* (0.0578)	-1.0503* (0.0651)	-1.3410* (0.1044)
lnG_{it}	0.6444* (0.0971)	1.3040* (0.1149)	2.0564* (0.3547)	1.5687* (0.1686)	2.9973* (0.1797)	2.2508* (0.2387)
Constant	1.2597* (0.2014)	2.9136* (0.2360)	2.0740* (0.4853)	3.2547* (0.3391)	6.2843* (0.2739)	4.2841* (0.6008)
F statistic	4 664.63	1 781.79	1 090.76	3 025.40	1 319.33	1 050.62
p-value	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
R^2	0.9708	-	-	0.9557	-	-
N	425	425	425	425	425	425
Groups	25	25	25	25	25	25
H$_0$: absence of autocorrelation in residuals of first order		p-value	0.008		p-value	0.011
H$_0$: absence of autocorrelation in residuals of second order		p-value	0.805		p-value	0.496
Hansen test		Prob>chi2	0.762			0.453
Sargan test		Prob>chi2	0.835			0.792

Source: Prepared by the authors.
Note: *denotes statistical significance at 1%.

V. Concluding remarks

This study has analysed the effects of economic growth and income inequality on poverty in Brazil between 1981 and 2013 using a dynamic panel data methodology and two measurements of poverty: the poverty and the extreme poverty headcount indices. Three periods were studied: the first runs from 1981 to 2013, while the second and third correspond to the stages prior to and following the implementation of the Real Plan (1981-1994 and 1995-2013, respectively).

The estimates arrived at using the OLS, within-group and system-GMM methods confirm the presence of persistent poverty dynamics and provide support for the statement that this phenomenon is primarily associated with the poor distribution of income in Brazilian society.

An analysis of the situation during the years leading up to the Real Plan (1981-1994) indicates that the income and inequality elasticities of poverty, as measured by both the poverty and the extreme poverty headcount indices, were lower (in absolute terms) than they were in the other two periods that were studied. This is probably due to the instability observed in the Brazilian economy in the 1980s. The combination of high inflation and a slow pace of growth in income during that decade can be identified as the reason why poverty levels became less sensitive to changes in income and inequality levels.

The analysis of the income elasticity of poverty indicates that growth was pro-poor in 1981-2013 according to the definition employed by Kakwani and Pernia (2000) of that term, but only in the case of extreme poverty. This means that income increased more than poverty declined. When the definition developed by Ravallion and Chen (2003) is used, then growth was pro-poor in all the periods that were studied as measured both by the poverty and the extreme poverty headcount indices.

Although, regardless of the definition used, growth was not pro-poor during the subperiod preceding the Real Plan, it was indeed pro-poor for both the poor and the extremely poor segments of the population in the subperiod that followed the implementation of the Plan. This result can be attributed to the economic stability seen during those years, the real increase in the minimum wage and the implementation of direct income transfer policies (França, 2010; Kakwani, Neri and Son, 2010).

The values obtained when the income and inequality elasticities of poverty were estimated indicate that policies aimed at boosting mean per capita household income and policies designed to reduce inequality tend to lower both poverty and extreme poverty levels. The results also indicate that these policies are more effective in reducing the extreme poverty headcount than the poverty headcount because, since the extreme poverty line is lower than the poverty line, it is more sensitive to policy measures.

A decline in inequality has a greater effect in terms of reductions in the extreme poverty and poverty headcounts than policies focused solely on boosting mean income do. In other words, according to the results presented here, economic growth policies that promote both an increase in income and a reduction in inequality (for example, policies targeting the poorest segments of the Brazilian population, such as the Bolsa Família programme) tend to reduce poverty more than economic growth policies whose sole aim is to raise mean income levels.

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The performance of the computer software and services sector in Argentina: microeconometric evidence on public-sector support programmes

Mariano Pereira, María Florencia Barletta and Gabriel Yoguel

Abstract

This article analyses the impact of public-sector support programmes on the recent performance of the computer software and services sector in Argentina. First, the effect of these programmes on firms' innovation performance is studied, with a propensity score matching technique used to calculate the average treatment effect on treated firms. The results confirm that receiving public funds had a positive impact on the ratio between research and development (R&D) spending and sales, employment in R&D and the propensity to introduce new products or processes. The effect of policy intervention on firms' economic performance is then analysed, with an instrumental variables design being used in this case. The results show a positive impact on the propensity to export, export intensity and employment growth.

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I.   Introduction

In the early 2000s, the outlook for the computer software and services sector in Argentina seemed extremely negative. Nobody foresaw that just over a decade later, employment in the sector would outstrip that in the automotive complex (by 8%) and in several primary sectors (agricultural services), industrial sectors (capital goods, leather and footwear, wood, furniture, iron and steel) and service sectors (electricity, gas and water), and would almost match that in sectors employing far more people in the late 1990s (banking, insurance and real-estate services). A number of published studies made it clear that the problems facing the sector limited its prospects for development. These constraints centred on three areas: (i) commercial capabilities that were too focused on the domestic market, which hampered the potential for developing an industry that could compete externally (López and Ramos, 2008); (ii) workers' limited technological capabilities, which constrained the level of complexity attainable in products and services (Borello and others, 2005), and (iii) the lack of a critical mass of prominent firms with in-house product development capabilities, which limited the scope for identifying successful sectoral profiles (Chudnovsky, López and Melitsko, 2001; López, 2003; Perazzo and others, 1999). Despite these constraints, though, the sector took a very different path. Since the devaluation of the Argentine peso in 2002, computer software and services firms have grown apace, with aggregate employment, sales and exports all rising much faster than in manufacturing industry (Barletta and others, 2013; Maldonado, Morero and Borrastero, 2013).

The main factors behind this performance have been: the increased competitiveness of the Argentine economy since the devaluation of early 2002, greater outsourcing of software development around the world (which dynamized a major segment of computer software and services firms) and other cultural and contextual characteristics favourable to participation in external markets (knowledge of English, time zones and availability of skilled human resources, among others). These factors were supplemented by some local ones, such as strong growth in the domestic market and the high level of capabilities generated in earlier decades at older firms and since the early 2000s in the great majority of the new ones, and by a large range of public programmes to promote quality certification, exports and research and development (R&D), which propelled innovation efforts. Over the past few years, a number of studies have looked at the determinants of the recent performance of the computer software and services sector in Argentina. However, surprisingly few have analysed the effects of innovation support programmes on the performance of the knowledge-intensive services sector in general and computer software and services in particular (Castro and Jorrat, 2013).

The present article has three specific aims: (i) to evaluate the causal effect of innovation support programmes on firms' innovation and economic performance; (ii) to analyse the existence of spillovers or indirect effects by considering links between beneficiary and non-beneficiary firms forged for the purpose of joint R&D, and (iii) to evaluate complementarity between instruments, since firms can apply for and receive support from more than one programme.

In accordance with these aims, the present article seeks to answer a number of questions. What role has been played by public-sector innovation support programmes in the innovation behaviour of computer software and services firms? Is it possible to say that sectoral incentive instruments have helped improve firms' economic performance via this more virtuous innovation behaviour? What role, ultimately, has been played by public-sector science and technology policy in the sector's strong performance over the last decade? These questions are answered using a database of 187 computer software and services firms constructed specially to capture the specificities of the sector. Two identification techniques are applied to this database to capture the average treatment effect on the treated among firms in receipt of subsidies, fiscal incentives or both in 2010: (i) propensity score matching (Rosenbaum and Rubin, 1983), and (ii) doubly robust estimations (Tsiatis, 2006; Leon, Tsiatis and Davidian, 2003; Lunceford and Davidian, 2004).

The article is divided into five sections, including this Introduction. The second section presents the database and the descriptive statistics for the differences between beneficiary and non-beneficiary firms. The third section describes the methodology proposed for capturing the causal effect of receipt of a public-sector benefit. The fourth estimates the causal effect of participating in a public-sector programme on a firm's innovation behaviour and economic performance. The fifth and last section contains the main conclusions of the article.

II.   The database

The database used for this article was compiled from a survey of firms in the computer software and services sector in 2011. The study received financing from the Carolina Foundation in Spain and assistance with sample design from the Employment and Business Dynamics Observatory of the Ministry of Labour, Employment and Social Security. The sample comprises a total of 189 firms, 2 of which were discarded because their data were incomplete. On average, the firms examined employed about 60 people each and declared annual sales of about 8 million pesos in 2010, with 56% stating they engaged in some foreign trade. Most of the firms in the sample began trading in the late 1990s, are Argentine in origin (93%) and operate independently, with just 10% forming part of a business group (see table 1).

Table 1
Argentina: descriptive statistics for computer software and services firms, 2010

	Observations	Mean	Standard deviation	Minimum	Maximum
Employment (*number of employees*)	187	58	14	1	1 500
Sales (*millions of pesos*)	187	7.9	22.7	0	200
Exports (0 = does not export; 1 = exports)	187	0.56	0.5	0	1
Starting year	187	1999	7	1968	2010
Part of business group (0 = not part; 1 = part)	187	0.1	0.3	0	1
Foreign ownership (*percentages*)	187	7.02	2.38	0	100

Source: Prepared by the authors.

To identify treatment status (i.e., whether a firm participated in some public-sector innovation support programme), the database contains a total of four binary variables. The first three indicate receipt of a benefit from the Argentine Technology Fund (FONTAR), the Trust Fund for the Promotion of the Software Industry (FONSOFT) and the Software Industry Promotion Act, respectively, while the fourth indicates participation in any of the three programmes (BENEF). The frequency distribution (see table 2) reveals that 23% of firms stated they had received a FONTAR benefit in 2010, 49% that they had participated in FONSOFT and 35% that they were registered for the purposes of the Software Industry Promotion Act. Overall, across all instruments, 65% received some kind of research, development and innovation (R&D&I) benefit.

Table 2
Argentina: participation by computer software and services firms in public programmes, 2010
(Percentages)

	FONTAR	FONSOFT	Software Industry Promotion Act	BENEF
Non-beneficiaries	77.25	51.32	64.55	34.39
Beneficiaries	22.75	48.68	35.45	65.61
Total	100	100	100	100

Source: Prepared by the authors.

The small sample size means that the possible differential effects of the various financing programmes on the innovation and economic behaviour of computer software and services firms cannot be robustly identified. Consequently, all that is estimated is the average treatment effect on the treated of participation in any of the sectoral support programmes (the BENEF variable).

As for control variables, the database makes it possible to design a heterogeneous set of variables that might affect the likelihood of innovation support being received. The number of employees in 2008 was taken as a proxy for firm size. Since the variable presents a skewed distribution, it was incorporated as a logarithm (l_size). The variable constructed to represent firms' age measures the number of years elapsing from the start of trading to 2008, and the bias in its distribution was corrected by employing the logarithmic form once again (l_age). Two variables were incorporated to complete the structural characterization of the sector: a dichotomous indicator for the firm's membership of a business group (ext_own) and another one, also dichotomous, to indicate whether the firm was located in Buenos Aires (bsas). As a proxy for the firm's capacity for absorption (Cohen and Levinthal, 1990), a binary variable indicating the presence of employees with postgraduate qualifications was created (cap_abs). Also incorporated were three variables for aspects of innovation behaviour within firms. The first is a binary variable indicating whether the firm conducts innovation efforts in the form of internal R&D (internal_r&d), the second captures the number of links established with other firms for R&D purposes (link_firms) and the third is the number of links with public institutions such as universities and technology centres declared by the firm (link_publ).[1] These latter variables can be expected to positively affect the likelihood of receiving a public-sector benefit. In other words, the expectation is, first, that firms making innovation efforts will be more active innovators and will be more likely to apply for and obtain public-sector funding, and, second, that firms forming part of networks (with other firms, universities or technology centres) to supplement their internal capabilities will be more active users of the instruments made available by the public sector to procure financing.

Lastly, the set of result variables encompasses the innovation and economic dimensions of firms' performance. The following four proxy variables are proposed for innovation performance: (i) a binary variable taking the value 1 if the firm introduced a new product or service into the market (inno); (ii) the ratio between the firm's investment in R&D and its sales (r&d_share); (iii) the ratio between the number of employees working in R&D and the total number of employees (emp_r&d), and (iv) the ratio between sales of the innovative product or service in dollars and the number of people working in the R&D department (r&d_product). In relation to firms' economic performance, three aspects are considered: (i) labour productivity, measured by sales per employee (product); (ii) export performance, measured using two indicators: a binary one indicating whether the firm exports (expo_bin) and a continuous one indicating the firm's export coefficient (expo_share), and (iii) employment performance, based on the cumulative rate of employment growth from 2008 to 2010 (emp_grow).

Table 3 presents the descriptive statistics for the sample variables. In the great majority of cases, the average value of the variables differs significantly between firms in the control group and those in the treatment group. For example, firms that received some support from the programmes analysed were characterized by more proactive innovation behaviour, manifested in higher indices of internal R&D, links with other firms and institutions and absorption capacity. As regards the set of result variables, although the test of means yields significant differences, these cannot be attributed to participation in the public programme. Systematic differences between the two groups relative to the group of covariates suggest that beneficiary firms are substantially different from their peers in the control group, so a direct comparison between them would be biased. The matching technique applied in this article allows a control group with observable characteristics similar to beneficiary firms' to be selected. After this exercise, the differences between the two groups can, subject to certain assumptions, be explained by participation in the public-sector support programme.

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[1]   To avoid spurious correlations, links with the Ministry of Industry and with the National Agency for the Promotion of Science and Technology (ANPCyT) of the Ministry of Science, Technology and Productive Innovation were not considered.

Table 3
Differences of means between treated and untreated firms

|	BENEF	
	Difference	*p*-value
Covariates		
l_size	0.31	0.13
l_age	0	0.83
ext_own	-3.24	0.38
bsas	0.11	0.10
cap_abs	0.06	0.67
internal_r&d	0.22	0.00
link_firms	0.12	0.35
link_publ	0.44	0.00
Innovation performance		
prod_inno	25.33	0.00
r&d_share	3.84	0.42
emp_r&d	0.04	0.78
r&d_produc	-72.904	0.33
Economic performance		
product	0.15	0.38
expo_bin	0.17	0.03
expo_share	0.26	0.00
emp_grow	0.01	0.09

Source: Prepared by the authors.

III. Methodological framework

1. An exploratory approach from the perspective of causal diagrams

The directed acyclic graphs (DAG) toolkit can be used to show the difficulties attendant on causal inference when the effects of participating in an innovation support programme are evaluated. Formally, a DAG (Pearl, 2000) is a graphic representation of the qualitative assumptions about causal relations that the researcher has in mind. More simply, it is a network composed of nodes (variables), arrows (suggested causal relations) and missing arrows (suggested non-existent causal relations). To this network may be added a set of assumptions and rules that serve to relate the graph rigorously to econometric data and models.

Figure 1 shows the DAG between a firm's (i) participation in a public-sector programme of R&D financing; (ii) innovation behaviour, and (iii) economic performance. In this diagram, each node represents a variable (or set of variables) with its corresponding distribution function. The arrow indicates the direction of the causal effect between two variables: if the arrow goes from innovation to performance, then what is being suggested is that the latter is causally affected by the former. The utility of this diagram is that it serves to illustrate the difficulties in identifying the causal effect associated with participation in a public-sector programme.

Figure 1
Acyclic diagram of causal relations
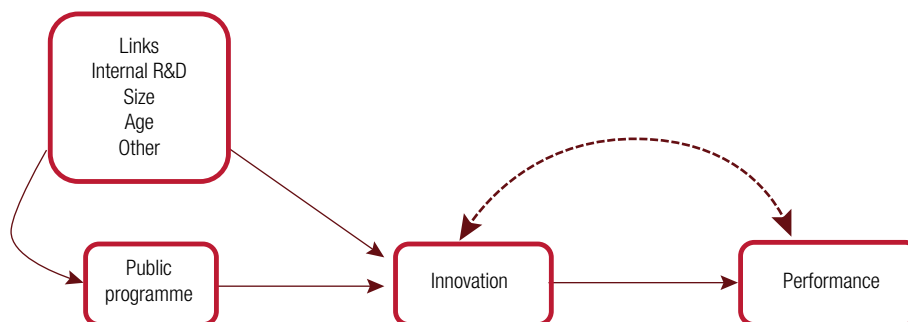
Source: Prepared by the authors.

First of all, we propose a causal relationship between receipt of financing for innovation activities (public programme) and a firm's innovation behaviour (innovation). The diagram indicates a set of covariates affecting both participation in the public programme and the introduction of new products or services. This group of covariates comprises observable factors such as firm size, age, ownership, geographical location, implementation of internal R&D activities, the presence of employees with postgraduate qualifications and links with other actors (firms, universities or public research centres). With this DAG, the use of an econometric model to estimate the relationship between the programme and innovation will produce a skewed and inconsistent result, since the causal effect may be "confused" with the impact of variations in the set of observable covariates. Consequently, two identification and estimation strategies are proposed to estimate the causal effect of the public programme: (i) a non-parametric estimate based on database balancing using the propensity score for participation in the public programme, and (ii) a doubly robust parametric estimation (Lunceford and Davidian, 2004).

Secondly, the diagram embodies the idea that the relationship between the public programme and a firm's economic performance is mediated by the latter's innovation performance. Consequently, the product of the effect of the innovation programme and the effect of innovation on performance should be used to estimate impact. Since these two variables are determined simultaneously by the presence of unobservable common factors, however (the dotted two-way arrow), the direct estimation proposed above will yield biased and inconsistent results. Thus, to identify this latter effect, an instrumental variables strategy needs to be adopted in which the group of instruments is formed by the set of covariates and participation in the public programme.

2. Identification strategy: propensity score matching

The impact evaluation literature offers a varied menu of techniques for estimating the causal effects of a public programme.[2] They include the following parametric and non-parametric estimation strategies in particular: (i) difference in differences; (ii) selection models; (iii) instrumental variables; (iv) propensity score matching, and (v) regression discontinuity.

In the present case, the structure of the database (presented and analysed in the next section) restricts this large menu. The lack of a cross-sectional and temporal database (micropanel) means that a difference in differences model cannot be estimated. Likewise, because no set of valid instruments is available to capture participation in the support programme, it is impossible to conduct a causal inference study based on instrumental variables methods. The same goes for the possibility of

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[2] See Heckman, Lalonde and Smith (1999), Imbens and Wooldridge (2009) and, for a broader approach, Morgan and Winship (2007).

constructing an exclusion equation to estimate a selection model. Thus, the cross-sectional character of the database means that only a propensity score matching technique can be employed.

To begin with, there is assumed to be a variable of interest for firm i, Y_i (the decision to innovate or R&D spending, for example), which can take two values depending on whether the firm participates in a public-sector financing programme. Then, D_i is defined as a binary variable taking the value 1 if firm i participated in the programme and 0 if it did not. The variable of interest can thus be defined as follows:

$$Y_i = \begin{cases} Y_{i0} \ if \ D_i = 0 \\ Y_{i1} \ if \ D_i = 1 \end{cases}$$

Similarly, the average causal effect of participating in the financing programme can be defined as follows:

$$E\left(\alpha_{TT}\right) = E\left[\left(Y_{i1} - Y_{i0}\right)\big/ D_i = 1\right] = E\left(Y_{i1}\big/ D_i = 1\right) - E\left(Y_{i0}\big/ D_i = 1\right)$$

When the causal effect of a policy is evaluated, responses are sought to a number of questions. What would have happened to the beneficiaries if the programme had not existed? What would have happened to R&D spending if the programme under examination had not been adopted? From an econometric perspective, this raises a missing data problem. Although we can directly observe $E\left(Y_{i1}\big/ D_i = 1\right)$ (the average value of the target variable among firms participating in the programme), this is not the case with $E\left(Y_{i0}\big/ D_i = 1\right)$ (the value the target variable would have had if the same firm i had not participated in the programme). It is only possible to see what happened, not what would have happened to the target variable without the programme. This hypothetical situation is known as a counterfactual. Since it cannot be observed, it must be estimated by employing the subgroup of firms that did not participate in the programme: $\widehat{Y_{i0}}$. Thus, the formula for calculating the average treatment effect on the firms participating in the programme is modified as follows:

$$E\left(\alpha_{TT}\right) = E\left[\left(Y_{i1} - \widehat{Y_{i0}}\right)\big/ D_i = 1\right]$$

Since the programme beneficiaries were not assigned randomly, it is not possible to rule out the possibility of selection bias when estimating the impact of D_i on Y_i. Consequently, it cannot be affirmed that $E\left(Y_{i0}\big/ D_i = 1\right) \neq E\left(Y_{i0}\big/ D_i = 0\right)$, nor can the counterfactual be estimated using the simple average for the firms that did not receive fiscal benefits or subsidies. To solve this problem, recourse may be had to the assumption of conditional independence (Rubin, 1977), according to which both participation in a public programme and the potential outcome of this are statistically independent for firms with the same set of observable characteristics, X. Thus, $E\left(Y_{i0}\big/ D_i = 1, X\right) = E\left(Y_{i0}\big/ D_i = 0, X\right)$ and the differences between the two groups are only attributable to the programme. Given these specifications, the average treatment effect can be estimated as:

$$E\left(\alpha_{TT}\right) = E\left(Y_{i1}\big/ D_i = 1, X = x\right) - E\left(Y_{i0}\big/ D_i = 0, X = x\right)$$

Some important points should be made. Besides the conditional independence assumption, another prerequisite for the consistency of the matching is for there to be adequate overlap between the control group and the group of firms receiving the subsidy. The control group needs to contain at least one unit that is similar enough to each treated firm. In practice, this condition is achieved by restricting the sample to a common support, for which purpose the lower and upper bounds for the propensity score are calculated and observations whose score lies outside these are discarded.

The exercise of matching each firm that participated in the programme introduces a further problem known as the "curse of dimensionality". The list of observable factors influencing both programme participation and the result studied may be too long, making it almost impossible to carry out separate matching for each individual unit. As the set of observable factors used during the matching procedure grows, the likelihood of finding an exact control drops exponentially. Again, it is very easy to prove in practice that using a relatively small set of factors to apply the matching method yields a set of beneficiaries for which there is no possible match. Rosenbaum and Rubin (1983) suggested matching beneficiaries and non-beneficiaries using their propensity score only (estimate of the conditional likelihood of participating in the programme). This reduces the matching procedure conceived as a multidimensional problem (where the dimension depends on the number of variables in the problem) to a one-dimensional problem. In practice, it is done by means of a maximum likelihood estimation for the probability of participating in the programme in the light of a set of covariates.

However, it is not enough to calculate the propensity score for each firm in the sample. Matching two firms with an identical score is almost impossible, since this is a continuous variable. Accordingly, the econometric literature documents the development of different methods for solving this problem. The present article uses kernel matching: each firm participating in the programme is matched against a weighted average of all the firms in the control group, with the weight they are given being inversely proportional to the distance between the propensity score of the treated firm and the firm in the control group.

Summed up, the basic idea of propensity score matching is to use statistical matching procedures to construct control groups. This methodology corrects for the observable differences between the treatment group (programme beneficiaries) and the control group (non-beneficiaries) by seeking out the individual unit in the non-beneficiary sample (which will form the control group) that best matches each individual unit in the treatment group sample. In this way, the differences between the two groups can only be attributed to participation in the public programme. The main assumption of this methodology is that participation is based on observable characteristics of the individual units. If this is not so, the evaluation results obtained with this methodology will be biased. The source of this distortion lies in the potential correlation between unobservable variables that affect both the individual's decision to participate in the programme and the variable of interest in the evaluation. Table 4 presents a schematic summary of the stages in this method.

Table 4
Steps in propensity score matching

Step 1	Specify and estimate a probit model for programme participation. Generate the probability predicted for each unit in the sample.
Step 2	Restrict the sample to a common support. All firms that have received some fiscal benefit or subsidy and that have a score above or below the maximum or minimum, respectively, are discarded from the control group.
Step 3	Estimate the average treatment effect on the treated using the kernel method with the following formula: $$\alpha_{TT}^{kernel} = \frac{1}{N^T} \sum_{i \in T} \left\{ Y_i^T - \frac{\sum_{j \in C} Y_j^C G\left(\frac{p_j - p_i}{h_n}\right)}{\sum_{k \in C} G\left(\frac{p_k - p_i}{h_n}\right)} \right\}$$ Where G(.) is a kernel function, h_n is a bandwidth parameter and $\dfrac{\sum_{j \in C} Y_j^C G\left(\frac{p_j - p_i}{h_n}\right)}{\sum_{k \in C} G\left(\frac{p_k - p_i}{h_n}\right)}$ is a consistent estimator of the counterfactual result.
Step 4	Apply bootstrapping to calculate the standard error associated with each average treatment effect on the treated and check whether the difference of means is statistically different from zero.

Source: Prepared by the authors, on the basis of D. Czarnitzki and C. Lopes-Bento, "Value for money? New microeconometric evidence on public R&D grants in Flanders", *Research Policy*, vol. 42, No. 1, Amsterdam, Elsevier, 2013.

3.   Identification strategy: doubly robust estimation

Robins and colleagues (Bang and Robins, 2005; Robins, Rotnitzky and Zhao, 1995; Robins, 2000) introduced the concept of doubly robust estimates. This technique requires both the model for estimating each firm's propensity score and the model for estimating the result variable to be used in the same estimator. The estimators calculated in this way are known as semi-parametric efficient estimators. Tsiatis (2006) and Leon, Tsiatis and Davidian (2003) show that these are doubly robust estimators because they yield consistent estimates of the average treatment effect on the treated, provided at least one of the following conditions is met: (i) the propensity score model is properly specified (i.e., $\hat{p}_i$ is the true propensity score), or (ii) the regression model relating the result variable with the covariates is properly specified. If both models are well specified, the estimates of the semi-parametric estimator will have the least variance. Lunceford and Davidian (2004) propose the following formula to calculate doubly robust estimates:

$$\hat{\tau}_{DR} = \frac{1}{N}\sum_{i=1}^{N}\frac{A_i Y_i - (A_i - \hat{p}_i)m_1(\underline{X}_i)}{\hat{p}_i} - \frac{1}{N}\sum_{i=1}^{N}\frac{(1 - A_i)Y_i + (A_i - \hat{p}_i)m_0(\underline{X}_i)}{1 - \hat{p}_i}$$

Where $\hat{p}_i$ is an estimate of the probability of participating in the public programme and $m_A = (\underline{X}_i) = E\left(Y_i \mid A_i = A, \underline{X}_i\right)$ for $A=0$ or $A=1$ are the predicted values derived from separate regressions of the result variable as a function of the set of covariates.

IV.   Evaluation of the causal effect of programme participation

1.   The effect on innovation behaviour: using propensity score matching to estimate the average treatment effect on the treated

As was emphasized in the methodology section, the first step in calculating the average treatment effect on the treated is to balance the sample between beneficiary and non-beneficiary firms on the basis of the observable factors. For this, it is necessary to estimate the probability of a computer software and services firm participating in one of the public programmes analysed in the light of its observable characteristics.

To match beneficiary and non-beneficiary firms, a probit model was used to estimate the probability of participation in some public programme. Table 5 presents the marginal effect of each of the variables proposed on the predicted probability. The joint significance test validates the model presented in all cases and suggests a fairly heterogeneous set of results that varies with the public programme analysed. For example, participation in FONTAR is positively affected by the firm's size in 2008, its internal R&D activities and its links both to other firms and to public institutions, including universities and technology centres. Where FONSOFT is concerned, the results indicate that the likelihood of receiving the benefit differs significantly by geographical location and is greater the more links the firm has to universities and technology centres, and the more internal R&D it conducts. The likelihood of receiving fiscal benefits under the Software Industry Promotion Act is positively affected by size and is greater when the firm has R&D links with other firms. Again, this likelihood is systematically higher for firms that conduct internal R&D and, paradoxically, lower for those that have staff with postgraduate qualifications. The last model analyses the likelihood of participating in any of

the programmes. There are found to be significant differences by geographical location, while internal R&D is associated with a positive impact, as are R&D links to other firms or public institutions.

Table 5
Marginal effects of the probit model on the participation dummy

	BENEF
l_size	0.015
l_age	2 676
ext_own	-0.001
bsas	0.155*
cap_abs	-0.025
internal_r&d	0.215**
link_firms	0.012**
link_publ	0.426***
Wald chi2(8)	43 295
Prob. > chi2	0.0000
No.	187

Source: Prepared by the authors.
Note: Heteroskedasticity-robust standard errors were calculated. *,** and *** indicate statistical significance at 10%, 5% and 1%, respectively.

As was pointed out earlier, an important prerequisite is for there to be a control observation with a quite similar propensity score for each firm receiving a benefit (common support restriction). For this, the minimum and maximum values of the propensity score were calculated for the firms in the control group, with the result that four firms in the treatment group that were not matched by similar firms in the control group were discarded and consequently not considered in the matching process.[3] Since the missing observations are not a substantial part of the sample, the common support restriction does not significantly affect the results. Besides the common support requirement, the position involved in the conditional independence assumption is that the differences between beneficiary and non-beneficiary firms, conditioned by the propensity score, are only attributable to the public programme being studied. From another angle, this assumption implies that differences between the two groups of firms based on observable factors disappear. Thus, the covariates proposed to explain participation in the public programme are required to be balanced between the two groups. Table 6 presents the *p*-value associated with the test of means and suggests that the null hypothesis that there are no differences between beneficiary and non-beneficiary firms cannot be rejected.

Table 6
Quality of matching: *p*-value of the test of means

	BENEF
l_size	0.686
l_age	0.78
ext_own	0.818
bsas	0.775
cap_abs	0.228
internal_r&d	0.712
link_firms	0.646
link_publ	0.782

Source: Prepared by the authors.

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[3] This only applies to FONTAR, since there were no units outside the common support region in the other programmes analysed.

Despite these results, it would be very unwise to assume that matching quality was good solely on the basis of the test of means.[4] Accordingly, and in line with Rosenbaum and Rubin (1983), the next step was to estimate the standardized average bias after matching to evaluate its quality. Standardized bias is defined as the difference of means in the two situations (treatment and non-treatment), divided by the square root of the mean of the respective variances. The consensus has set a limit of 10% to evaluate the quality of matching. Table 7 presents the percentage of bias for each covariate. The criterion is met as a rule, with some one-off exceptions.

Table 7
Quality of matching: percentage of standardized bias

	BENEF
l_size	-5.1
l_age	3.7
ext_own	2.3
bsas	-3.7
cap_abs	7.7
internal_r&d	-4.2
link_firms	-6.2
link_publ	3.2

Source: Prepared by the authors.

In conclusion, evaluation of the results of the test of means and the standardized bias values shows that all the covariates proposed are well balanced after matching. The matching can be pronounced successful and the average treatment effect on the treated can now be estimated. If there are statistically significant differences, these can be attributed to the benefit received.

It could be argued that the foregoing results lack robustness owing to the small number of beneficiary firms in the sample for some instruments. For this reason, a fourth estimation is proposed, in which the average treatment effect on the treated is calculated by considering participation in any of the public programmes mentioned. The calculations show that for firms in the computer software and services sector, receipt of some benefit was associated with R&D spending as a share of sales that was almost 6 percentage points higher on average than it would have been in the absence of State aid. The share of R&D personnel in firms' staffing totals also increased sharply with participation in public-sector support programmes, with the estimate showing a rise of 11 percentage points. Lastly, estimation of the average treatment effect on the treated revealed that the likelihood of introducing a new product into the market was an average of 20% greater in firms with benefits than in those without (see table 8).

Table 8
Estimates of the average treatment effect on the treated of programmes overall

	No. of beneficiaries	Effect (without matching)	Effect (with matching)
prod_inno	124	25.00	19.00**
r&d_share	124	5.28	5.45*
r&d_produc	124	-72.904	-48.601
emp_r&d	124	0.04	0.11**

Source: Prepared by the authors.
Note: Standard errors were calculated by bootstrapping with 300 repetitions. *,** and *** indicate statistical significance at 10%, 5% and 1%, respectively.

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[4] The test of means involves very strong assumptions, such as normal distribution of the variables (which is clearly not met when there are binary or multinomial variables), and is highly sensitive to the amount of data in the sample.

Taken all together, the results show that public-sector instruments supporting the sector helped to configure a group of firms with more dynamic innovation behaviour. In particular, higher R&D spending went to recruit staff focused on these activities. This set of activities explains why innovation is systematically more likely at firms that have received some public-sector benefit. Despite these substantive results, though, no significant differences were observed in the productivity of R&D teams (measured as sales of the innovative product per employee on the R&D team) following receipt of the subsidy. This outcome could be due to the shortness of the period analysed (2008 to 2010), with further time being needed for more dynamic innovation behaviour to translate into improved productivity in the area of R&D.

2. The effect on innovation behaviour: using doubly robust estimates to estimate the average treatment effect on the treated

This section applies an alternative technique to estimate the average effect of participating in a public-sector programme of support for R&D, known as doubly robust estimates.

First, two alterations are made to the group of variables considered. One is that the treatment variable takes the value 1 if the firm participates in any of the financing programmes studied (FONTAR, FONSOFT or the Software Industry Promotion Act). The other is that the group of result variables considered is restricted to: (i) R&D spending as a percentage of sales; (ii) personnel employed on R&D as a percentage of the total, and (iii) introduction of a new product on the market.

Doubly robust estimation is another parametric technique that combines matching based on the propensity score with an econometric model. The model proposes a function for each independent variable, including the covariates affecting the decision to participate in the relevant financing programme and a binary variable for participation in some public programme. The estimation is also restricted to observations belonging to the common support region, and each firm is weighted by its propensity score. The results presented in table 9 bear out the overall significance of the model proposed (p-value of 0%), the individual significance of the fiscal benefits or subsidies and the non-significance of the remaining variables. This set of results shows that the estimation of the average treatment effect on the treated is not biased and is consistent, as the selection bias based on observable characteristics is controlled for. It can be seen that the estimates obtained are similar to the results arrived at by non-parametric methods, as detailed in table 8.

Table 9
Doubly robust estimates of the average treatment effect on the treated

	R&D spending	Human resources in R&D	Product innovation
Benefit	6.003*	0.093**	0.212**
Size	-4 552	-0.167	-0.055
Age	483.09	-5.02	6 399
Internal R&D	-3.220	0.145	0.273
R&D links	-2 498	-0.038	0.013
Links with universities etc.	-4 428	-0.008	0.019
Postgraduate staff	6 549	-0.006	0.034
Group	0.003	0	0.001
Geographical location	-2 236	-0.058	-0.089
Intercept	-3 639 123	39 097	-48 149
Prob. > chi2	2.21E-07	0.0101238	0.0130124
No.	43	173	184

Source: Prepared by the authors.
Note: Robust standard errors were calculated. *,** and *** indicate statistical significance at 10%, 5% and 1%, respectively.

3. The effect on firms' economic performance: instrumental variables

The main purpose of this article is to determine the role of public instruments for R&D financing on the recent performance of the computer software and services sector. The methodological approach adopted was based on the idea that the effect on a firm's performance was mediated by innovation (see figure 1), and it was explained that because the introduction of a new product on the market was contemporaneous with the economic performance being measured, this causal effect was not being captured. The response proposed was to use an instrumental variables technique involving generation of a localized exogenous variation from variables that are redundant when it comes to explaining a firm's performance but highly correlated with the likelihood of innovation.

Three dimensions were taken into account in estimating firms' performance: (i) labour productivity (measured as sales per employee); (ii) export performance (two indicators are used: a binary one indicating whether the firm exports, and a continuous one indicating the firm's export coefficient), and (iii) employment performance (two variables are used: the total number of employees in 2010 and the cumulative growth rate since 2008). The set of observable factors and public programme participation are used to capture the innovation effect.

Table 10 presents the results of the second stage of the instrumental variables method. The maximum likelihood projection obtained in the previous stage (see table 10) was used as an innovation instrument. The statistical inference confirms that innovation, and thence public programme participation, positively affected firms' performance. In particular, the results show a causal effect on export performance (both the probability of selling to international markets and the export coefficient) and the rate of employment growth. Lastly, no causal effect on labour productivity is captured.

Table 10
The effect of innovation on firms' performance

	Labour productivity	Exports (binary)	Export coefficient	Employment growth
inno_prod	0.288	2.295***	1.770**	1.461*
l_size	0.172**	0.244***	0.170**	-0.139*
l_age	-17 296	9 727	25 563	69.928***
ext_own	-0.010**	0.010*	0.013***	0.017***
bsas	0.195	-0.137	-0.16	0.051
Intercept		-75 965		
chi2	12 227	20 228	27 154	39 186
P	0.0318	0.0011	0.0001	0.0000
No.	171	188	185	182

Source: Prepared by the authors.
Note: Robust standard errors were calculated. *,** and *** indicate statistical significance at 10%, 5% and 1%, respectively.

V.   Conclusions

The computer software and services sector in Argentina has performed remarkably well in recent years, becoming one of the most dynamic sectors in the country. Growth figures for domestic sales, exports and employment have been much higher than those for manufacturing and other types of services. Just 15 years ago it was unthinkable that by 2014 the computer software and services sector would be employing more staff than the automotive complex in the aggregate. The present document sets out

from this context to provide a general characterization of the sector with a view to understanding the factors that determined firms' economic and innovation performance. Broadly speaking, this growth is explained by a varied set of factors associated with the local and global economic context of the past 15 years. Among them, public policy in support of the sector, implemented by national funds such as FONTAR and FONSOFT and underpinned by the Software Industry Promotion Act of 2004, seems to have played an important role.

The descriptive statistics presented in this document show that over half the firms surveyed had access to some type of public-sector benefit. Comparing beneficiary firms with non-beneficiaries reveals that the former have a higher ratio of R&D investment to sales, more quality certifications and more links to other firms and institutions for the purpose of accessing outside knowledge to supplement their own capabilities. Thus, the evidence found supports the hypothesis that public-sector programmes in support of innovation have helped to configure a group of firms characterized by more dynamic innovation behaviour.

A number of econometric exercises were then carried out to evaluate specifically the role of public programmes in firms' economic and innovation performance. The results bring to light a chain of causal reactions that show, in the first instance, a positive effect accruing to firms from programmes supporting innovation behaviour. In some dimensions, such as the intensity of R&D spending (whether this is measured as the ratio between spending and sales or the percentage of workers dedicated exclusively to R&D work), participation in the public programme was followed by a large increase on the levels that would have been achieved in the absence of State intervention. In parallel with this, receipt of funds to supplement R&D investment was associated with more firms bringing a new product or service on to the market.

It was then shown that, taking the effect on the likelihood of innovation as a yardstick, public programmes in support of the sector were also a determinant of firms' performance in the market. This causal effect proved significant for firms' performance in international markets and in terms of employment, but not for productivity.

Consequently, it can be said that vertical promotion instruments targeted at the computer software and services sector played an important role in its rise after the country's fixed exchange-rate regime ended. Public intervention, and particularly industrial and technological policy, is a determinant that is often underestimated in the literature by comparison with exogenous factors, such as competitiveness gains deriving from devaluation, or favourable cultural factors, such as mastery of English or the time zone. This document provides quantitative information which it is hoped will contribute to a greater appreciation of the role of public intervention, particularly when this is conducted with vertically designed instruments like those analysed in this article.

Methodologically, it is important to signal two prior constraints on the impact study that need to be taken into account when the results are evaluated. First, the sample was constructed not for the explicit purpose of conducting an impact assessment, but to analyse the extent to which the technological, organizational and absorption capacity of computer software and services firms and their links to other firms, universities and public institutions supporting the sector explained their economic and innovation performance in recent years. Consequently, the variables used to capture treatment status (e.g., whether or not the firm participated in any of the programmes analysed) did not fully capture this phenomenon. For example, it is not possible to specify the year in which firms accessed these funds (all that is known is that they participated in the programmes between 2008 and 2010), or what the type of instrument was (non-refundable aid, fiscal credit or subsidized credit). Secondly, the cross-sectional nature of the database ruled out the use of evaluation techniques that could yield more precise results, such as a difference in differences study. With these limitations in view, though, the impact assessment design proposed the employment of all techniques that could

feasibly be used: regression with controls, propensity score matching and doubly robust estimates. The results of the estimations were fairly similar across the different techniques, which is evidence for their robustness. In any event, a very important recommendation regarding the availability of databases for implementing policy evaluation exercises is prompted by these limitations. That is, there is a need for databases designed so that they measure result variables at different points in time and can be cross-matched with information from the register of beneficiary firms, for the treatment and control groups to be specified more accurately.

Lastly, this study raises a number of questions of relevance to the agenda for future research on the role of support instruments in the computer software and services sector. First, there is a need for further analysis of firms accessing more than one innovation support programme. The results presented indicate that programme simultaneity does not affect firms' innovation performance. This suggests a need for greater complementarity between the bodies implementing public policy in order to improve efficiency and avoid incentives overlapping. Since the firms applying for and receiving more than one benefit at a time are likely to be the best performers, there should be an assessment of how far differentiated types of stimulus are required as firms grow. Second, after 15 years' continuous growth in the computer software and services sector, it seems important to have a debate on its future strategic orientation, i.e., to work out a way of specializing in the segments that have been most successful (finance, security and big data, among others) and support the creation of a critical mass of firms that can achieve a more complex specialization profile. Public instruments should be redesigned in the light of this debate with a view to targeting support at these strategic segments.

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Annex A1

This section briefly summarizes the programmes analysed in the article.

(a) The Software Industry Promotion Act

In 2004, the Argentine Congress passed the Software Industry Promotion Act (Law No. 25922). This law gave tax breaks to computer software and services firms and created a trust fund called FONSOFT (presented below). For a firm to receive the fiscal benefits established by the law, its main activity must be the creation, design, development, and production and implementation of software systems and associated technical documentation. The software must be wholly developed in Argentina.

Firms also need to do two of the following three things: (i) carry out their own R&D spending; (ii) hold quality certification, and (iii) export.

The fiscal benefits consist in:

(i)   Fiscal stability while the incentive regime is in force (until the end of 2019), with the authorities being debarred from altering the tax system for the sector operated in by firms enrolled in the incentive regime.

(ii)   An implicit reduction in social security charges: 70% of these charges can be credited via payment of other national taxes (except corporation tax).

(iii)   Reduction of up to 60% in the corporation tax payable.

(b) FONSOFT

As mentioned above, the Trust Fund for the Promotion of the Software Industry (FONSOFT) was created in 2004 under the Software Industry Promotion Act. FONSOFT is a programme that uses subsidies and credits to finance R&D spending by firms in the sector and start-up costs for firms entering it. The programme is involved at all stages of innovation projects, from formulation and implementation to evaluation of the results. This financing is only available to small and medium-sized enterprises (SMEs) dedicated to the production of software goods and services.

FONSOFT is organized around four instruments: (i) subsidies to firms to improve their quality standards or R&D spending; (ii) credits to firms to help them begin exporting or consolidate their export activities; (iii) subsidies to business owners to set up new firms, and (iv) matching grants to finance human resources training.

(c) FONTAR

The Argentine Technology Fund (FONTAR) is another of the funds run by the National Agency for the Promotion of Science and Technology (ANPCyT).[5] The beneficiaries of FONTAR are innovative firms (particularly SMEs) and public-sector organizations providing technical assistance to the private sector. Funding covers implementation of innovation projects in the following ways: (i) support for

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[5]   ANPCyT is part of the Ministry of Science, Technology and Productive Innovation and was set up in 1996 with the goal of promoting and supporting science and technology.

implementation of R&D projects; (ii) technical support for firms looking to draw up projects; (iii) technical and financial advice for firms requiring funding; (iv) financing of the projects selected, and (v) issuance of a certificate of conformity for R&D teams or departments in accordance with Law No. 23877.

The programme is involved at all stages of the innovation project, from formulation and implementation to evaluation of the results. FONTAR is organized into components, subcomponents and incentive instruments. In practice, firms apply for a particular instrument. Each has its own selection criteria and provides support at different stages of the innovation process.

FONTAR comprises two components: technological innovation in the production sector, and integrated products in production clusters. The former is further divided into three subcomponents that include: (i) technological modernization projects financed with subsidies and loans; (ii) support for technological development at SMEs (including the preparation of patent applications) via subsidies, and (iii) capacity-building for the provision of services to the production sector via refundable loans to non-profit organizations and private institutions.

Structural changes in Brazilian industry (1995-2009)

Helena Loiola de Figueiredo and
Maria Aparecida Silva Oliveira

Abstract

This article analyses the structural changes that took place in Brazilian industry between 1995 and 2009, by considering their intersectoral relations, through input-output analysis using the structural decomposition method and the calculation of linkage indices. The results show that the expansion of final demand plays a key role in industry growth in terms of employment, value added and gross production value. Natural-resource-intensive industry has grown particularly strongly. Another finding is that intersectoral demand has weakened, particularly in scale-intensive sectors that use differentiated technology.

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I.   Introduction

Achieving economic development in a country requires overcoming a series of obstacles, and implementing structural change to ensure that the economy's resources flow rapidly towards modern economic activities of higher productivity (Rodrik, 2013). Many recent studies focus on the different structural changes that have taken place in the Brazilian economy in recent years, focusing particularly on those that occurred in the productive sectors of the economy following the trade liberalization process of the 1990s.

Each study examines fundamental aspects that enrich that debate, such as the industrial sector's declining share in the economy,[1] the spread of outsourcing,[2] Dutch disease[3] and the regressive specialization of the industrial sector,[4] among others. Although there is no consensus on this matter, authors who emphasize the importance of industry for economic dynamics agree on the peculiarities of this sector. Hirschman (1958) claims that industry has more forward and backward linkages than the agriculture and service sectors, and that positive externalities and indirect effects are likely to be more prominent in that sector. Consequently, industry growth would have greater positive effects on the economy as a whole. Kaldor (1966) also stresses the differentiated role of industry with respect to technology and the greater potential of static and dynamic economies of scale.

With the aim of contributing to that debate, this paper seeks to analyse the structural changes that took place in Brazilian industry between 1995 and 2009, by considering its intersectoral relations. Following the trade liberalization process, Brazilian industry entered a period of structural specialization, which was premature in terms of the country's per capita income (Carvalho and Kupfer, 2011). The liberalization that took place in the 1990s favoured sectors of the economy that were already consolidated, and it led to a reduction in the industry share in value added and employment, the spread of outsourcing and the growth of the services sector. Such circumstances can imply a shift in industrial structure towards sectors of lesser technological content.[5] The study of the type of specialization occurring in Brazilian industry is an important element in the design of general economic policy.

To analyse structural change in Brazilian industry, this article uses the structural decomposition method of input-output analysis proposed by Miller and Blair (2009). This has been used in studies of both national and international scope to analyse structural changes both in the industrial sector and in the economy as a whole. Franke and Kalmbach (2005) studied structural change in industry and its effects on the service sector in Germany in the 1990s. Linden and Dietzenbacher (2000) analysed the determinants of structural change in the European Union by applying the RAS method to update the coefficients of an input-output matrix that captures variations of substitution and manufacturing effect. Guilhoto and others (2001a) compared the structural change that took place in Brazil between 1959 and 1980 with that occurring in the United States between 1958 and 1977; and they split the change in the structure of the economies into three components, inside and outside the sector (final demand, technology and intersectoral relations).

In addition to the structural decomposition, this article analyses the trend of linkages between industrial sectors according to the Rasmussen-Hirschman and Ghosh linkage indices; and it deflates

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[1] See Oreiro and Feijó (2010), Marquetti (2002) and Bonelli (2005).

[2] See Araújo (2010).

[3] See Bresser-Pereira (2009) and Bresser-Pereira and Marconi (2008).

[4] See Carvalho and Kupfer (2007 and 2008).

[5] See Shafaeddin (2005).

the data from the matrices as proposed by Miller and Blair (2009).[6] The data are obtained from input-output tables based on a classification of 42 sectors,[7] estimated by Guilhoto and Sesso Filho (2010) on the basis of the national accounts produced by the Brazilian Geographical and Statistical Institute (IBGE), to analyse changes in production value, value added and employment in Brazilian industry between 1995 and 2009. The choice of period was based on criteria of data availability and quality, and comparability with other studies.

The importance of evaluating structural change in Brazilian industry is easy to justify. The pursuit of economic development requires taking account of several aspects of industrialization for policy definition. As argued by Rodrik (2007), productive specialization based on comparative advantages, particularly in commodities, is unlikely to be sufficient to achieve industrial upgrading. For that process to take place, investments are needed in other large manufacturing sectors, such as machinery, equipment and production facilities. The analysis of the type of specialization taking place in national industry is useful for evaluating public industrial development policies. This paper classifies the industrial sectors by technological level.[8]

In view of the above, this article is divided into five sections including the Introduction. Section II describes the methodology used in the study, while section III considers the source and classification of the data considered. Section IV presents the study's main findings, and section V summarizes the conclusions.

II.   Methodology

This section firstly describes how the input-output matrices were deflated. The base year was 2009, and the method used, defined in Miller and Blair (2009), is referred to as double deflation, since it consists in a two-stage process. The first step involves deflating intermediate demand, final demand and gross production value (GPV), using a price index for each sector calculated based on data from the tables of uses and resources published in the IBGE national accounts. The price index by sector, with base year 1995 = 100 for the 42 sectors and the period 1996-2009, was:

$$I_{x,1995} = 100, \ \text{for} \ x = sec \, tor \, 1, \dots, sec \, tor \, 42 \tag{1}$$

$$I_{x,1996} = I_{x,1995} \cdot annual \ variation \ in \ prices \ of \ sector \ x \ _{1996}$$

$$I_{x,t} = I_{x,t-1} \cdot annual \ variation \ of \ prices \ of \ sector \ x_t, \ for \ t = 1997, \dots, 2009 \tag{2}$$

The indices were re-based to 2009 as follows:

$$I_{t,base \ year \ 2009} = \left(I_{t,base \ year \ 1995} \Big/ I_{2009, base \ year \ 1995}\right) \cdot 100, \ \text{for} \ t = 1995, \dots, 2009 \tag{3}$$

Once calculated, a vector π_t could be constructed with the price indices for the 42 sectors, and it was possible to deflate intermediate demand (Z^b), final demand (f^b) and GPV (X^b):

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[6] As input-output data from different years are being compared, it is important to distinguish changes attributed to prices from those of other sources.

[7]   The 42 sectors are presented in the annex.

[8] The sectors were classified by technological intensity based on Organization for Economic Cooperation and Development (OECD) (2005), which draws on Pavitt (1984). See the annex.

$$\pi_t = \begin{bmatrix} I_{agriculture,\ t} \\ I_{extractive\ mining,\ t} \\ \vdots \\ I_{non\text{-}commercial\ private\ services,\ t} \end{bmatrix}, \text{ for } t = 1995,\ldots,2009 \tag{4}$$

$$Z^b = \hat{\pi}_t Z_t \tag{5}$$

$$f^b = \hat{\pi}_t f_t \tag{6}$$

$$X^b = \hat{\pi}_t X_t \tag{7}$$

The second step consists in calculating the price index to deflate the value-added data. All data are obtained at current prices, including value added. The value added (v^b) needed to ensure that the GPV is the same in both the sum of the rows and the sum of the columns, is calculated as follows:

$$\left(v^b\right)' = \left(X^b\right)' - i' Z^b \tag{8}$$

The deflator of value-added can then be calculated as,

$$\hat{r}_t = \hat{v}^b \left(\hat{v}_t\right)^{-1}, \text{ where } v_t \text{ represents value added at current prices} \tag{9}$$

According to Miller and Blair (2009), although the double deflation method is widely used, it has many disadvantages for deflating input-output tables, since all the elements of the row of the transactions matrix are deflated by the same index. The authors point out that inter-industry prices can vary considerably in many economies, so deflation by the same index could be incorrect. An alternative method is the bi-proportional adjustment algorithm or RAS. The double deflation method was chosen, because the RAS method is used mostly to update and project coefficients, rather than to deflate input-output matrices.[9]

After the deflation, the linkage indices and structural decompositions are calculated. For this purpose, a number of input-output analysis relations need to be defined. An economy is assumed to consist of n sectors; X is an $n \times 1$ vector of sectoral gross production values; A is an $n \times n$ matrix of technical coefficients; and f is an $n \times 1$ vector of final demand for the output of each sector. The sector production vector X can be expressed through the equation $X = AX + f$. The necessary algebraic manipulations give the input-output model that relates the respective sectoral outputs.

$$X = Lf \tag{10}$$

where $L = (I-A)^{-1}$, is an identity matrix of order n. $(I-A)^{-1}$ is the matrix of technical coefficients of direct and indirect inputs, or the Leontief inverse matrix, which captures the direct and indirect effects of exogenous changes in final demand in the production of the n sectors.

There are various ways of measuring sectoral linkages. This paper calculates the Rasmussen-Hirschman backward linkage indices, created by Rasmussen (1956) and subsequently developed

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[9] According to Miller and Blair (2009), the RAS technique requires less information and is widely used to estimate input output matrices that are not available. Using an input output matrix for a given year $\left(A_0\right)$ the RAS technique makes it possible to estimate the matrix for a later year $\left(\check{A}_1\right)$, knowing the sum of the rows $\left(\sum_{j=1}^{n} Z_{1j}\right)$, the columns $\left(\sum_{j=1}^{n} Z_{i1}\right)$ and the gross production value of all sectors of the economy in the later year.

by Hirschman (1958). According to Guilhoto and Sesso Filho (2010), the backward linkage indices indicate how much each sector demands from the other sectors of the economy. The indices in question are based on the inverse Leontief matrix ($L=(I-A)^{-1}$), so l_{ij} can be defined as an element of the matrix L, along with L^* which is the mean of all elements of L. It is also possible to calculate L_{*j}, which is the sum of the elements of a column of L, with n representing the number of sectors in the economy. Algebraically this gives:

$$L_{*j} = \sum_{j=1}^{n} l_{ij} \quad i,j = 1, 2, \ldots \tag{11}$$

Thus it is possible to determine the backward linkage indices:

$$U_j = \left[\frac{L_{*j}}{n} \right] \Big/ L^* \tag{12}$$

To calculate forward linkage indices, the Ghosh model is the most suitable (Miller and Blair, 2009). These indices show how a sector is demanded by the other sectors, or how it supplies them with inputs. Instead of considering the technical coefficient ($a_{ij}=Z_{ij}/X_j$), this model considers the coefficient of allocation of production ($b_{ij}=Z_{ij}/X_i$). The indices are based on the inverse Ghosh matrix ($G=(I-K)^{-1}$), so that g_{ij} can be defined as an element of the matrix G, along with G^* which is the mean of all elements of G. In addition, G_{i*}, can be calculated as the sum of the elements of a row of G. Algebraically, this gives:

$$G_{i*} = \sum_{i=1}^{n} g_{ij} \quad i,j = 1, 2, \ldots \tag{13}$$

It is thus possible to determine the forward linkage indices:

$$U_i = \left[\frac{G_{i*}}{n} \right] \Big/ G^* \tag{14}$$

The key sectors will be those that simultaneously show forward and backward linkage indices with values above one. Table 1 classifies the sectors by the value of their indices (less than or greater than one).

Table 1
Classification of intersectoral linkages

		Total forward linkage	
		Low (<1)	High (>1)
Total backward linkage	Low (<1)	(I) Independent	(II) Dependent on intersectoral demand
	High (>1)	(III) Dependent on intersectoral supply	(IV) Generally dependent (or key sector)

Source: Prepared by the authors, on the basis of R.E. Miller and P.D. Blair, *Input-Output Analysis: Foundations and Extensions*, New Jersey, Prentice-Hall, 2009.

Next, a structural decomposition is performed on GPV (ΔX), industrial employment $\Delta \varepsilon$ and value-added in industry (ΔV), assuming that there are input-output matrices for two periods (0 and 1). Then, the GPV for the two periods is obtained from equation (10), as follows:

$$X_0 = L_0 f_0; \ X_1 = L_1 f_1 \tag{15}$$

where, f_t is the final demand vector in year t; and $L_t = (I - A_t)^{-1}$ is the Leontief impact matrix in year t. The variation in GPV between the two years is:

$$\Delta X = X_1 - X_0 = L_1 f_1 - L_0 f_0 \qquad (16)$$

The structural decomposition method involves several comparative statics exercises in which various coefficients are changed to be able to compare the activity levels with a benchmark (Miernyk, 1974). Considering equations:

$$\Delta f = f_1 - f_0; \quad f_1 = (f_0 + \Delta f); \quad f_0 = (f_1 - \Delta f);$$
$$\Delta L = L_1 - L_0; \quad L_1 = (L_0 + \Delta L); \quad L_0 = (L_1 + \Delta L)$$

and substituting in equation (16) gives:

$$\Delta X = L_1(f_0 + \Delta f) - (L_1 + \Delta L)f_0 = (\Delta L)f_0 + L_1(\Delta f) \qquad (17)$$

The first part of equation (17) concerns technological change, whereas the second reflects changes in final demand. Although several combinations are possible, Miller and Blair (2009) perform the decomposition additively and develop some of those examples. Apart from that shown in equation (17) there are the following:

$$\Delta X = (\Delta L)f_0 + L_0(\Delta f) - (\Delta L)(\Delta f) \qquad (18)$$

$$\Delta X = (\Delta L)f_1 + L_1(\Delta f) - (\Delta L)(\Delta f) \qquad (19)$$

While all of the above equations are possible, Dietzenbacher and Los (1998) found that the combination of equations (18) and (19) is the most appropriate. That combination gives rise to equation (20), which is used in this study.

$$\Delta X = \left(\frac{1}{2}\right)(\Delta L)(f_0 + f_1) + \left(\frac{1}{2}\right)(L_0 + L_1)(\Delta f) \qquad (20)$$

The first term on the right-hand side represents the variation in GPV if there is a change in technology (it assumes the change in the inverse Leontief matrix - ΔL), whereas the second term captures the effect of variations in final demand (Δf) on ΔX.

As the calculation is based on the changes in the Leontief matrix, the effect of the technological change shows how the linkages between the sectors vary (weakening or strengthening of the link). The factors explaining the technological changes are: innovations, import substitution, an increase in the benefits obtained from economies of scale, changes in product mix (with the adoption of new substitutes or complimentary inputs in the production process), variation in relative prices (given that the technical coefficients in the Leontief matrix arise from the monetary valuation), and changes in trade patterns (exports and also import substitution). Those factors alter the technical coefficients in the Leontief matrix, and are shown in the calculated effect of the technological changes (Schuschny, 2005).[10]

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[10] Although structural decomposition makes it possible to identify the activities that recorded increases in output owing to technological change, the model does not contain information to identify and analyse its causes. In other words, this method cannot determine the increase in output of a given sector owing to the variation in each separate factor that comprises the technological change (innovation, economies of scale, changes in product mix, variation in relative prices, changes in trade patterns).

For the decomposition of employment, $(e_t)' = [e_{0,t} \ldots e_{1,t}]$ is the vector of employment coefficients representing the quantity of labour per monetary unit of production in sector i during period t. The inverse of those coefficient represents an indirect measurement of labour productivity, defined as:

$$e_{i,t} = \varepsilon_{i,t} \big/ X_{i,t} \tag{21}$$

Thus, the sectoral employment vector in period t will be:

$$\varepsilon_t = \hat{e}_t X_t = \hat{e}_t L_t f_t \tag{22}$$

and the vector of changes in employment will be:

$$\Delta \varepsilon = \varepsilon_1 - \varepsilon_0 = \hat{e}_1 L_1 f_1 - \hat{e}_0 L_0 f_0 \tag{23}$$

Using the same relations as in the decomposition of production, equation (23) can be written as follows:

$$\Delta \varepsilon = \left(\frac{1}{2}\right)(\Delta \hat{e})(L_1 f_1 + L_0 f_0) + \left(\frac{1}{2}\right)\left[\hat{e}_0 \Delta L f_1 + \hat{e}_1 \Delta L f_0\right] + \left(\frac{1}{2}\right)(\hat{e}_0 L_0 + \hat{e}_1 L_1)(\Delta f) \tag{24}$$

The first term of equation (24) is the portion of the variation in employment caused by changes in the direct labour coefficient. The second term represents the portion of the variation in sectoral employment owing to the technological changes that altered the input requirements of the production activities. The third term captures the effect of the variation in final demand on sectoral employment.

Lastly, the decomposition of value added is similar to that of employment. The difference is that account is taken of the direct vector of coefficients of value added, which is represented by the ratio between value added and output value ($va_{i,t}$), instead of using the vector of direct employment coefficients.

$$va_{i,t} = V_{i,t} \big/ X_{i,t} \tag{25}$$

$$V_t = \widehat{va_t} X_t = \widehat{va_t} L_t f_t \tag{26}$$

$$\Delta V = V_1 - V_0 = \widehat{va_1} L_1 f_1 - \widehat{va_0} L_0 f_0 \tag{27}$$

$$\Delta V = \left(\frac{1}{2}\right)(\Delta \widehat{va})(L_1 f_1 + L_0 f_0) + \left(\frac{1}{2}\right)\left[\widehat{va_0} \Delta L f_1 + \widehat{va_1} \Delta L f_0\right] + \left(\frac{1}{2}\right)(\widehat{va_0} L_0 + \widehat{va_1} L_1)(\Delta f) \tag{28}$$

III. Source and classification of the data

Input-output tables were used, based on a classification of 42 sectors, estimated by Guilhoto and Sesso Filho (2010) and available online at the website of the University of São Paulo Regional and Urban Economics Lab (NEREUS).

Thirty of the 42 sectors examined in this paper are industrial. The data published by the cited source adhere to an international classification criterion based on version 1.0 of the National

Classification of Economic Activities (CNAE).[11] The classification by technological intensity was done using the methodology based on the taxonomy created by Pavitt (1984) and adopted by the Organization for Economic Cooperation and Development (OECD) in various studies (OECD, 1987 and 2005), which was also used by Nassif (2006) to analyse Brazil's foreign trade, among numerous other studies.

Lall (2000) used that taxonomy in an analysis of technological change and industrialization in Asia. For that author, the main factor in the competitiveness of natural-resource-intensive sectors is access to the natural resources itself, whereas the competitiveness of labour-intensive sectors depends on the availability of low- and medium-skilled labour at a low relative cost compared with other countries. Scale-intensive sectors are those in which it is possible to gain by producing on a large scale. In sectors with differentiated technology, the products satisfy different patterns of demand; while the key competitive factor in sectors with science-based technology is rapid application of the science to the industrial technology.

A shortcoming of the Pavitt (1984) taxonomy is that the classification does not capture any of the changes that have occurred in the world economy over the last three decades. According to Dupas (1998), corporate production and distribution strategies have been reformulated, and the vertically integrated enterprise has given way to networks which incorporate different firms in a single global project. In this process, technology and capital become increasingly mobile, driven by the potential fragmentation of production chains. In this context, the oil and gas sector, which, in that taxonomy is classified as natural-resource-intensive, involves knowledge relating to oil extraction and refining, which would be technology-intensive. The same situation applies to the electronic appliances sector, classified as an industry of differentiated technology, in which the production chain includes segments with very similar characteristics to those of labour-intensive sectors. This article reconciles the sectors classified in that taxonomy by OECD (2005) and the sectors of the input-output matrix.[12]

IV. Results

1. Intersectoral relations

Use of the Ghosh and Rasmussen-Hirschman indices made it possible to ascertain how the production structure of the Brazilian economy has changed through time. This is shown in figures 1 and 2, which resemble encephalograms in medicine, measuring differences based on a certain standard. That analogy, originally defined in Guilhoto and others (2001a and 2001b), has been called the electro-econogram of the production structure. It can be constructed by showing the results of the indices on a linear graph. The smaller the change reflected in the dispersion of the graphs, the closer the results will be to those of the first year of the analysis (1995). In contrast, the larger the dispersion of the graphs, the larger will be the changes recorded in the productive structure over the period analysed.

The visual analysis shows that the Ghosh forward linkage indices (see figure 1) display decreasing variation, which indicates less intensity in intersectoral supply in the later years of the series, compared to the start of the period. That loss of intensity occurred slowly between 1995 and 2009, so the dispersion of the graphs was less than in the analysis of Rasmussen-Hirschman the backward linkage index (see figure 2). In this case, there is a clear change in the pattern of the graphs: slow change in the production structure in 1995-2000, followed by a clear change as from 2001, indicating a reduction in intersectoral demand as from that year.

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[11] The 30 industrial sectors form part of section C (extractive industries) and section D (manufacturing industries) of CNAE 1.0.

[12] See annex.

Figure 1
Electro-econogram of the Ghosh forward linkage index of the sectors of Brazilian industry, 1995-2009
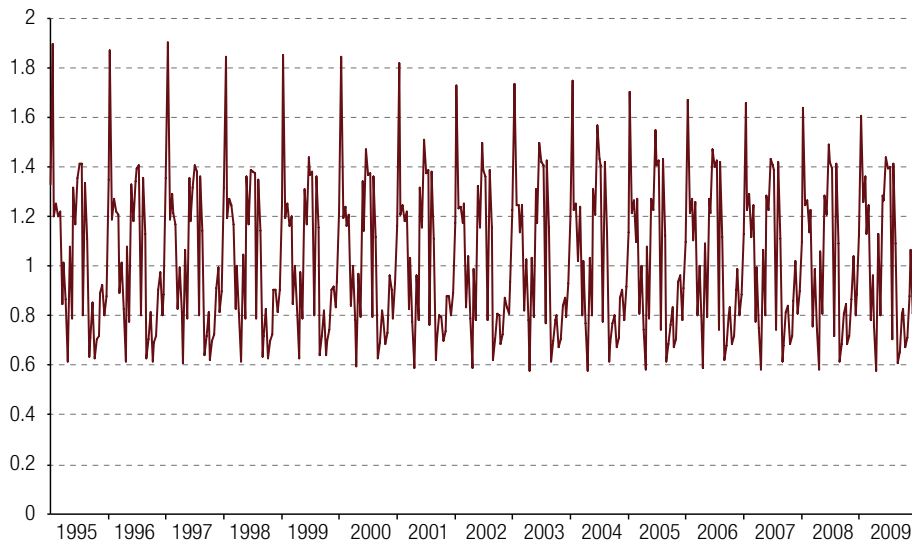
Source: Prepared by the authors, on the basis of data from the University of São Paulo Regional and Urban Economics Lab (NEREUS).

Figure 2
Electro-econogram of the Rasmussen-Hirschman backward linkage index of the sectors of Brazilian industry, 1995-2009
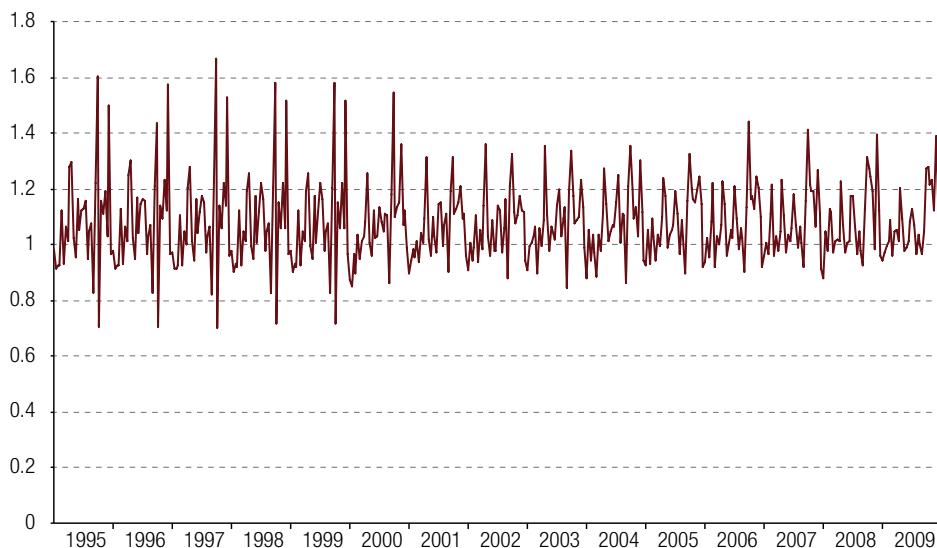
Source: Prepared by the authors, on the basis of data from the University of São Paulo Regional and Urban Economics Lab (NEREUS).

A factor that may have influenced this phenomenon was the change in the macroeconomic policy regime from 1999 onwards. Between 1995 and 1998, the macroeconomic policy regime based on an exchange rate anchor was implemented, which underpinned the success of the Real Plan. In 1999, that regime was replaced by the "macroeconomic tripod," which was in force until 2005, consisting of inflation targets, primary surplus targets, and a relatively free-floating nominal exchange rate. The change since 2001 is probably related to an increase in import content; in other words, the impact of foreign trade on the structure of national production, owing to the change in the economic

development model being pursued (particularly by the floating exchange rate as from 1999).[13] Although the integration of the Brazilian market into the global economy in the first decade of the twenty-first century made it possible to increase exports, many firms replaced national suppliers with foreign ones.[14] Thus, the reduction in sectoral demand relations (see figure 2) can be explained in terms of substitution by imported inputs.

Sectoral relations decline in both cases, since there was a reduction in sectoral relations in terms of input supply (captured by the Ghosh forward linkage index) and a much larger loss in sectoral demand relations (captured by the Rasmussen-Hirschman backward linkage index).

There follows a more detailed analysis of the sector linkages, in order to determine whether the intersectoral linkages were strengthened or weakened. The aim is to establish which sectors increase their linkages with the others, within the composition of manufacturing industry. Table 2 shows the backward and forward linkage indices.

Table 2
Rasmussen-Hirschman backward linkage indices and Ghosh forward linkage indices
1995 and 2009

Classification by type of technology	Sector	Ranking (highest backward linkage indices)		Backward linkage index		Ranking (highest forward linkage indices)		Forward linkage index	
		1995	2009	1995	2009	1995	2009	1995	2009
Science-based	Pharmaceutical and veterinary	24th	26th	0.95	0.97	22nd	23rd	0.81	0.8
Natural-resource-intensive	Extractive mineral	21st	30th	0.98	0.94	5th	5th	1.35	1.36
	Oil and gas	28th	25th	0.91	0.97	1st	1st	1.87	1.9
	Non-metallic mineral	27th	22nd	0.92	1	11th	10th	1.19	1.19
	Oil refining	10th	7th	1.13	1.13	3rd	2nd	1.39	1.41
	Coffee industry	1st	3rd	1.6	1.27	21st	20th	0.82	0.82
	Processing of vegetable products	30th	2nd	0.71	1.28	29th	29th	0.62	0.62
	Animal slaughter	8th	5th	1.16	1.21	27th	27th	0.69	0.7
	Dairy products industry	13th	4th	1.11	1.23	25th	25th	0.72	0.72
	Sugar manufacture	6th	8th	1.19	1.12	17th	17th	0.91	0.91
	Manufacture of vegetable oils	18th	1st	1.03	1.39	16th	15th	0.98	1
	Other products food	2nd	9th	1.5	1.12	23rd	21st	0.8	0.82
	Non-ferrous metallurgy	12th	11th	1.12	1.09	9th	9th	1.22	1.22
Labour-intensive	Other metallurgy	25th	28th	0.93	0.96	10th	12th	1.21	1.17
	Wood and furniture	23rd	24th	0.95	0.98	24th	24th	0.78	0.79
	Textile industry	14th	21st	1.07	0.99	13th	13th	1.13	1.14
	Garments	29th	27th	0.82	0.96	28th	28th	0.63	0.64
	Footwear manufacture	5th	15th	1.22	1.05	26th	26th	0.7	0.72
	Miscellaneous industries	22nd	29th	0.96	0.96	19th	18th	0.89	0.9

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[13] That argument is at the heart of studies on Dutch disease by Bresser-Pereira (2009) and Oreiro and Feijó (2010).

[14] In his study, Magacho (2013) decomposed changes in industry between 1995 and 2008 to determine the sectors in which substitution by imported inputs was most intense. The results show that in the primary sectors the impact of replacement by imported inputs in output was 13.9%. In the high and medium-high technology sectors, replacement by imported inputs reduced the growth of output value by 18.1%, particularly in the chemical and electrical equipment sectors. The author states that the national process of replacement by imported inputs was most intense between 2003 and 2008.

Table 2 (concluded)

Classification by type of technology	Sector	Ranking (highest backward linkage indices)		Backward linkage index		Ranking (highest forward linkage indices)		Forward linkage index	
		1995	2009	1995	2009	1995	2009	1995	2009
Scale-intensive	Iron and steel	26th	20th	0.93	1.01	8th	8th	1.27	1.29
	Automobiles trucks and buses	3rd	6th	1.3	1.21	30th	30th	0.61	0.61
	Vehicle parts	19th	10th	1.02	1.09	14th	14th	1.08	1.07
	Cellulose, paper and graphics	7th	23rd	1.16	0.99	7th	6th	1.33	1.36
	Rubber industry	16th	18th	1.05	1.01	12th	11th	1.18	1.18
	Chemical elements	11th	12th	1.12	1.08	6th	7th	1.34	1.32
	Miscellaneous chemicals	9th	13th	1.16	1.07	2nd	3rd	1.41	1.38
	Plastics	17th	17th	1.05	1.04	4th	4th	1.35	1.36
Differentiated	Machinery and equipment	15th	16th	1.06	1.05	18th	19th	0.89	0.83
	Electrical material	20th	14th	1.01	1.05	15th	16th	1.01	0.99
	Electronic equipment	4th	19th	1.28	1.01	20th	22nd	0.83	0.81

Source: Prepared by the authors, on the basis of data from the University of São Paulo Regional and Urban Economics Lab (NEREUS).

Comparing 1995 with 2009, there is evidence that the backward linkages that increased most were those of the sectors in the natural-resource-intensive group, in other words their intersectoral demand increased. Whereas in 1995, just two of the five highest indices of backward linkage corresponded to natural-resource-intensive sectors (coffee industry, other food products), by 2009 that number had risen to five (manufacture of vegetable oils, processing of vegetable products, coffee industry, dairy product industry, and animal slaughtering). In addition, another four sectors in that group raised their positions in the ranking: oil and gas, non-metallic minerals, oil refining, and non-ferrous metallurgy. In the forward linkages of those sectors, there were not many changes between the two years, except in the case of the manufacturer of vegetable oils, which became a key sector in 2009, because its forward linkage index rose by more than one. Oil refining and non-ferrous metallurgy were key sectors in both 1995 and 2009.

In general, the group of labour-intensive sectors lost positions among the highest indices of backward linkage between 1995 and 2009, except for the clothing sector. In 1995, the textile industry and footwear manufacture displayed backward linkage indices above one, and were dynamic sectors from the standpoint of intersectoral supply.[15] In 2009, the textile industry had lost backward linkages and was no longer a dynamic sector from the supply standpoint. Moreover, although the forward sector linkages of those sectors increased, their positions among the highest forward linkage indices remained unchanged. The textile industry was the only sector in the group to record forward linkage indices greater than one in both 1995 and 2009, which means that it is dynamic from the demand perspective. The loss of backward linkages in that sector caused a change in classification from a key sector in 1995, to a sector dependent on intersectoral demand.

Among scale-intensive sectors, there has been an increase in backward sector linkages in the iron and steel industries and vehicle parts, which meant that iron and steel was classified as a key sector in 2009. The other sectors in the group suffered reductions. All sectors display backward linkage indicators above one, so they are dynamic sectors from the intersectoral supply viewpoint. Only the pulp, paper and graphics sector ceased to be dynamic in terms of supply in 2009, owing to the loss

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[15] A sector that is dynamic from the supply standpoint succeeds in influencing intersectoral supply, since it is an important source of demand in other sectors of the economy. In contrast, a dynamic sector from the demand standpoint is an important supplier to other sectors and influences intersectoral demand.

of backward linkages, and thus was not classified as a key sector in that year. Moreover, despite the forward linkages remaining broadly unchanged between 1995 and 2009, the corresponding indices were greater than one in nearly all sectors of the group, which shows that they are dynamic from the demand standpoint. The automobile, trucks and bus industry is an exception, because it displayed a forward linkage index less than one in 1995, and in 2009 was not classified as a key sector because it was dependent on intersectoral supply and because it provides input to the other sectors of the economy. The scale-intensive group had the largest number of key sectors in the period.

In the pharmaceutical and veterinary sector, which uses science-based technologies, there were few changes in the backward and forward linkages. That sector was dynamic neither from the supply nor the demand standpoint, because its backward and forward linkage indices were less than one in both 1995 and 2009. Its backward linkages increased and forward linkages declined in the period analysed.

Lastly, all sectors with differentiated technology displayed backward linkage indices greater than one in 1995 and 2009, and thus represented dynamic sectors in terms of supply. Between 1995 and 2009, the backward linkages decreased in the machinery and equipment and electronic equipment sectors, and increased in the electrical material sector. In terms of forward linkages, only the electrical material sector recorded an index greater than one in 1995. Nonetheless, as it lost linkages, it ceased to be dynamic from the intersectoral demand standpoint, and was reclassified as dependent on intersectoral supply.

While there were no significant changes in forward linkages between 1995 and 2009, the situation is different when the backward linkages are analysed. In general, there was a loss of backward linkages in the Brazilian economy, which implies a weakening of intersectoral demand. The analysis of technological groups shows that while the natural-resource-intensive sectors gain importance in 2009, by increasing the demand for inputs from the other sectors (displaying higher backward linkage indices in that year), the backward linkages declined in important scale-intensive demanding sectors,[16] along with differentiated technology[17] and two labour-intensive sectors (textile industry and footwear manufacture).

The analysis of sector linkages shows that the Brazilian economy went through a process in which the intersectoral demand of natural-resource-intensive sectors increased by more than the sectors of the other groups. There was an industrial restructuring process in which the scope of the intermediate demand of natural-resource-intensive sectors increased, while that of the scale-intensive sectors and those with differentiated technology declined. The results of the structural decomposition of employment, value-added and GPV are presented next.

2. Decomposition of employment

Table 3 summarizes the results of the decomposition of employment in the industrial sectors by type of technology used. Of the 2.7 million jobs created between 1995 and 2009, 931,200 were in sectors that use natural resources-intensive technologies, and 951,980 were in labour-intensive sectors. This concurs with the claim by Nassif (2006) that sectors with natural resource-based and labour-intensive technologies have greater capacity to generate direct jobs.

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[16] Except for iron and steel and vehicle parts.

[17] Except for machinery and equipment.

Classification of the industry by type of technology	Total variation in employment	Contribution of technological change		Contribution of change in final demand		Contribution of the change in the direct labour coefficient	
Science-based	30.24	-27.58	-91.22%	148.60	491.40%	-90.77	-300.18%
Natural-resource-intensive	931.10	-22.34	-2.40%	1 627.80	174.82%	-674.36	-72.43%
Labour-intensive	951.98	-361.43	-37.97%	802.99	84.35%	510.43	53.62%
Scale-intensive	466.08	-28.96	-6.21%	1 078.39	231.37%	-583.35	-125.16%
Differentiated	376.68	-113.58	-30.15%	527.49	140.04%	-37.24	-9.89%
Total	2 756.09	-553.90	-20.10%	4 185.28	151.86%	-875.29	-31.76%

Source: Prepared by the authors, on the basis of data from the University of São Paulo Regional and Urban Economics Lab (NEREUS).

The largest increases in employment in the period analysed took place in the sectors of vegetable product processing (360,000 jobs created) and garments (357,000 jobs). As the first of these is natural-resource-intensive and the second is labour-intensive, their contributions were important for the results of those groups.

Technological change resulted in job losses in the industry, particularly in labour-intensive sectors (361,430 fewer jobs). The garment sector made a big contribution to that result, because it was the manufacturing industry sector that suffered the largest job loss for that reason. Despite the negative technological effect, the second largest increase in employment occurred in that same sector. This was because the contributions of the change in final demand and direct labour coefficient together outweighed the contribution made by technological change. The same happened in the other sectors of the labour-intensive group.

Final demand increased industrial employment, particularly in natural-resource-intensive sectors (creation of 1.62 million jobs) and in scale-intensive sectors (1.07 million new jobs). The sectors recording the largest increases in employment as a result of final demand were, respectively, other food products, and pulp, paper and graphics.

The direct labour coefficient reduced employment in all of the groups analysed, except for the labour-intensive group. The coefficient measures the amount of labour per monetary unit of production, and its inverse can be interpreted as an indirect measure of labour productivity. Thus, the largest increase in labour productivity (with a reduction of 674,360 jobs owing to the change in the coefficient) occurred in the natural-resource-intensive sectors. In the labour-intensive sectors, labour productivity declined (with an increase of 510,430 jobs), because the garment and footwear manufacturing sectors lost productivity during the period analysed.

The restructuring of industrial employment between 1995 and 2009 occurred with the largest job growth in the technology and natural-resource-intensive sectors. Those sectors suffered a reduction owing to the technological effect, but, while the labour-intensive sectors lost productivity, the natural-resource-intensive sectors displayed the largest increase in employment owing to increases in final demand and labour productivity in the period analysed.

3. Decomposition of value added

In terms of the decomposition of value added, the industrial sectors with natural-resource-intensive technology recorded the largest share in the total variation of industry value-added (41.07%), followed by the scale-intensive sectors (35.26%), sectors using differentiated technology (15.40%), science-based sectors (5.62%) and, lastly, the labour-intensive sectors (2.65%) (see table 4).

Table 4
Brazil: structural decomposition of industrial value added by type of technology, 1995-2009
(Millions of reais at 2009 prices and percentages)

Industry classification by type of technology	Total variation in value-added		Variation in value added attributable to each effect					
			Technological change		Change in final demand		Change in the direct coefficient of value-added	
Industry	161 915.17	100.00%	-7 103.70	100.00%	199 620.03	100.00%	-30 601.15	100.00%
Science-based	9 099.46	5.62%	-2 653.79	37.36%	14 706.55	7.37%	-2 953.30	9.65%
Natural-resource-intensive	66 496.22	41.07%	12 767.17	-179.73%	76 979.45	38.56%	-23 250.40	75.98%
Labour-intensive	4 288.48	2.65%	-7 977.85	112.31%	22 625.39	11.33%	-10 359.06	33.85%
Scale-intensive	57 088.73	35.26%	-3 804.19	53.55%	60 025.71	30.07%	867.21	-2.83%
Differentiated	24 942.29	15.40%	-5 435.04	76.51%	25 282.93	12.67%	5 094.39	-16.65%

Source: Prepared by the authors, on the basis of data from the University of São Paulo Regional and Urban Economics Lab (NEREUS) and the national accounts published by the Brazilian Geographical and Statistical Institute (IBGE).

The technological changes that occurred in the Brazilian economy and the period examined, contributed to the increase in value-added only in the natural-resource-intensive sectors, but declined in the other sectors (see table 4). Thus, the contribution of the effect of technological changes to the change in value added in the different groups was negative in the industrial sectors that use science-based technology (-29.16%), and also in labour-intensive sectors (-186.03%), scale-intensive sectors (-6.66%), and sectors using differentiated technologies (-21.79%). Only in the sectors using natural-resource-intensive technology was an increase in value added owing to technological change, which accounted for 19.20% of the variation in value added of that group (see table 5). The contribution made by the effect of the change in final demand was the factor explaining most of the growth in value added in all of the groups analysed. In particular, the change in final demand is the only factor explaining the increase in value added of the labour-intensive sectors, since the contributions made by the other factors were negative (see table 5).

Table 5
Brazil: contributions made by the different the effects to the variation in value added, 1995-2009
(Percentages)

Industry classification by type of technology	Contributions of the different effects		
	Technological change	Change in final demand	Change in the direct coefficient of value-added
Industry	-4.39%	123.29%	-18.90%
Science-based	-29.16%	161.62%	-32.46%
Natural-resource-intensive	19.20%	115.77%	-34.96%
Labour-intensive	-186.03%	527.59%	-241.56%
Scale-intensive	-6.66%	105.14%	1.52%
Differentiated	-21.79%	101.37%	20.42%

Source: Prepared by the authors, on the basis of data from the University of São Paulo Regional and Urban Economics Lab (NEREUS) and the national accounts published by the Brazilian Geographical and Statistical Institute (IBGE).

The effect of the direct coefficient of value added implied reductions in all of the groups except for the sectors with scale-intensive and differentiated technology. In other words, the ratio between value added and production value increased only in those two groups. This result shows that the sectors that require skilled labour, such as sectors using scale-intensive and differentiated technology, are those that succeeded in increasing their capacity to generate greater income or value added for the

economy. In contrast, the sectors that carry out standardized or codified tasks, requiring little skilled labour, do not achieve that result, as is the case with several natural resource- and labour-intensive sectors. In the maquila industries[18] for example, the ratio between value added and production value is small.

The analysis of value added reveals a process of re-primarization or backward specialization in Brazilian manufacturing industry, which is reflected in the greater share of sectors with natural-resource-intensive technology in the total variation in industry value added. Backward specialization means a change in the composition of industry, such that natural-resource-intensive sectors become relatively more important.

In an analysis of "deindustrialization" from the standpoint of value added at the sector level between 2000 and 2008, Morceiro (2012) obtained interesting results. Most of the sectors of manufacturing industry analysed by the author recorded a positive performance. The largest reductions occurred in the sectors of electronics material and communications equipment, clothing and accessories, wood products and leather articles and footwear. According to that author (p. 109):

> Thus, although deindustrialization (real downsizing) proceeded in those four sectors, it can be considered localized or concentrated in a few sectors that are more exposed to international competition, particularly Asian (characterized by countries that are strong in the labour-intensive sectors and in the electronics industry). Nonetheless, other labour-intensive sectors such as the textiles, furniture, articles of rubber and plastic, and basic metallurgy, performed poorly compared to the manufacturing industry; and, if no measures adopted, they could be susceptible to deindustrialization.

The sectors identified by Morceiro (2012) as being exposed to international competition, especially Asian, are labour-intensive sectors (such as wood and furniture, clothing and footwear manufacture), and those with differentiated technology (such as electronics equipment). The only one of those four sectors in which value added decreased between 1995 and 2009 was the garment industry, with a contraction of R$ 14,143.35 million at 2009 prices, owing to the negative contribution of final demand. The sector was responsible for the low share of the labour-intensive group in the total variation in industry value added (see table 4).

In the period under analysis, value added also declined in the vegetable product processing sectors (- R$ 15,781.08 million at 2009 prices) and vegetable oil manufacture (- R$ 9,413.70 million at 2009 prices), both of which are natural-resource-intensive sectors. Despite the reduction in those sectors, the substantial results of the oil and gas industries (variation of R$ 22,001.17 million in value added) and oil refining (total variation of R$ 20,159.69 million in value added), meant that the natural-resource-intensive sectors generated the largest share of the total variation of industry value added (see table 4).

The results obtained partially confirm the study by Morceiro (2012). It was found that labour-intensive sectors were those that participated least in the growth of industry value added in the period analysed, because they suffer from international competition. The discrepancies observed are due to methodological differences between the studies. The structural decomposition was used to analyse the total variation in value added between 1995 and 2009, and the factors explaining that variation (technological changes, changes in final demand, and changes in the direct coefficient of value-added). Morceiro (2012) compared the growth rates of value added of 2000 and 2008 by industry sector, and used a different industrial sector classification than that used in this paper.

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[18] The maquila industries were originally labour-intensive manufacturing enterprises set up on the border between Mexico and the United States to exploit cheap labour. The production process that was transferred to Mexico was very simple and did not require skilled labour, nor did it contribute to technological change. See Bresser-Pereira, Marconi and Oreiro (2009). As those industries also involve product assembly based on imported components, they generate little value added.

4. Decomposition of gross production value

Table 6 shows the results of the decomposition of industry GPV by type of technology. Between 1995 and 2009, industrial GPV increased by R$ 722,743.92 million at 2009 prices. The largest share in the total increase in industry GPV corresponds to the natural-resource-intensive sectors (51.08%), followed by the scale-intensive sectors (31.28%), and then those with differentiated technology (8.75%), labour-intensive sectors (5.05%), and those using science-based technology (3.84%).

Technological change in the economy reduced industry GPV by R$ 12,770.66 million, and the sectors using differentiated technology were those that contributed in most to that result. In addition, technological change reduced industry GPV in all groups analysed, except for natural-resource-intensive sectors, where GPV increased by R$ 47,447.81 million as a result of technological change (see table 6).

Table 6
Brazil: structural decomposition of industrial gross production value (GPV) by type
of technology, 1995-2009
(Millions of reais at 2009 prices and percentages)

Industry classification by type of technology	Total variation in GPV		Variation in GPV owing to the effects			
			Technological change		Change in final demand	
Industry	722 743.92	100.00%	-12 770.66	100.00%	735 514.57	100.00%
Science-based	27 771.54	3.84%	-6 010.14	47.06%	33 781.68	4.59%
Natural-resource-intensive	369 201.13	51.08%	47 624.88	-372.92%	321 576.25	43.72%
Labour-intensive	36 474.48	5.05%	-17 116.81	134.03%	53 591.29	7.29%
Scale-intensive	226 064.67	31.28%	-12 580.79	98.51%	238 645.46	32.45%
Differentiated	63 232.10	8.75%	-24 687.80	193.32%	87 919.89	11.95%

Source: Prepared by the authors, on the basis of data from the University of São Paulo Regional and Urban Economics Lab (NEREUS) and the national accounts published by the Brazilian Geographical and Statistical Institute (IBGE).

The contribution made by the effect of technological change to the variation in value added of the different groups was negative for the industry sectors that use science-based technology (-21.64%), and also for labour-intensive sectors (-46.93%), scale-intensive sectors (-5.57%) and those using differentiated technology (-39.04%). It was positive only for the natural-resource-intensive sectors (12.90%) (see table 7).

Table 7
Brazil: contribution of the different effects to the variation in gross production value
(GPV), 1995-2009
(Percentages)

Industry classification by type of technology	Contribution of the effects	
	Technological change	Change in final demand
Industry	-1.77%	101.77%
Science-based	-21.64%	121.64%
Natural-resource-intensive	12.90%	87.10%
Labour-intensive	-46.93%	146.93%
Scale-intensive	-5.57%	105.57%
Differentiated	-39.04%	139.04%

Source: Prepared by the authors, on the basis of data from the University of São Paulo Regional and Urban Economics Lab (NEREUS) and the national accounts published by the Brazilian Geographical and Statistical Institute (IBGE).

The technical changes that occurred in the Brazilian economy increased value added and GPV only in the natural-resource-intensive industrial sectors, whereas the other sectors suffered reductions. Those changes could reflect innovations, import substitution, economies of scale, changes in product mix, variations in relative prices, and changes in trade patterns.

Growth of final demand was the main factor driving an increase in industry GPV. Although that increase was greater in the natural-resource-intensive sectors (see table 6), the labour-intensive sectors were those displaying the largest contribution of the effect of the change in final demand (146.93% —see table 7). It was the impulse generated by the increase in final demand that caused science-based, labour-intensive, scale-intensive and differentiated technology sectors to record GPV growth in 1995-2009.

The analyses of industry GPV and value added confirm the hypothesis of a process of industrial re-primarization or backward specialization, because the largest increase in GPV corresponded to natural-resource-intensive sectors, followed by the scale-intensive sectors, those using differentiated technology, science-based technology, and labour-intensive sectors. The sectors that contributed most to the increase in GPV in the natural-resource-intensive group were oil refining and oil and gas.

The second largest increase in GPV occurred in the scale-intensive sectors, thanks to the contribution of the automobile, truck and bus sector, which had the largest increase in GPV within this group. In sectors using differentiated technology, the machinery and equipment sector achieved the highest GPV growth, which, by contrast, declined in the electronic equipment sector. The latter was affected by external competition, mainly from Asia. GPV growth in the labour-intensive sector was less than in the natural-resource-intensive sectors, and also in the scale-intensive and differentiated-technology sectors, owing to the reduction in GPV in the clothing sector, which is also affected by external competition.

V. Conclusions and final thoughts

This study analysed the structural changes that occurred in Brazilian industry between 1995 and 2009, considering its intersectoral relations. Empirical evidence showed that the structural changes weakened intersectoral demand in Brazilian industry. It was also found that an industrial restructuring process took place, as the scope of the intersectoral demand expanded in natural-resource-intensive sectors, but shrank in the scale-intensive and differentiated-technology sectors.

The restructuring of industrial employment between 1995 and 2009 occurred with the largest increases in the number of jobs in technology-intensive and natural-resource-intensive sectors. Although those sectors suffered a reduction owing to the technological effect, they displayed the largest increase in employment owing to growth in final demand in the period analysed.

The analyses of value added and GPV confirm the hypothesis of a process of industry re-primarization and backward specialization, because the largest increases in value-added and GPV corresponded to the natural-resource-intensive group.

An important conclusion of the study was that the technological changes made a negative contribution to employment growth, value added and GPV in Brazilian industry. The technological changes that took place in the economy contributed to reductions of 20.10% in employment, 4.39% in value added and 1.77% in GPV of Brazilian industry between 1995 and 2009. The labour-intensive sectors were the most affected from the standpoint of loss of jobs and value added, while sectors using differentiated technology with the most affected with respect to the reduction in GPV. Those changes could be due to technological innovations, an increase in the benefits obtained from economies of scale, changes in product mix, variation in relative prices or changes in trade patterns. One of the

technological changes that seems to be most important is the replacement of national inputs with imported ones, since this could be one of the causes of the reduction in the sectors' backward linkage indices, reflecting a weakening of the intersectoral demand of industry. In contrast, the contributions of final demand were responsible for most of the positive results of the structural decompositions.

In short, while the expansion of final demand played an important role in industry growth in terms of employment, value added and GPV, there was also a weakening of intersectoral demand in Brazilian industry. This was characterized by an increase in imported inputs for production, while that production is being financed above all by the growth of final demand.

Moreover, the importance gained by the natural-resource-intensive sectors, from the growth of their intersectoral demands or significant increases in employment, value added and GPV, do not seem sufficient for industry to promote dynamism in the Brazilian economy as argued by Rodrik (2007).

If public policy-makers want to balance the contributions made by the increase in final demand with stimulus for technological change to promote industry growth, incentives need to be provided for those changes to take place. In this context, industrial policy becomes the key public policy mechanism for creating favourable conditions for the survival of the industrial sector, mainly for the sectors with greater technological intensity.

Over the last decade, the government put the topic of industrial development policy back on the public policy agenda, and sought to implement an industrial policy, despite various obstacles, such as changes in international conditions after the crisis, the overvaluation of the real and the sudden emergence of China as a global economic force.

The economic liberalization of the 1990s and the complex process of productive and financial globalization decisively influenced business and corporate strategies and are still doing so today. That process seriously compromises government's capacity to develop national policies to strengthen industrial competitiveness.

The challenge is that this context requires the State to play a central role in mobilizing and coordinating the productive, technological, financial and organizational-institutional resources needed to make investments viable. In addition to guaranteeing coordination capacity, industrial policy must take account of the specifics of reality in its various dimensions (sectoral, technological, financial, organizational, institutional and regional), based on a long-term dynamic perspective.

This study contributes to the analysis of the sectoral dimension of industry by showing the trend of those sectors in terms of employment, value added, GPV and linkages. Nonetheless, the research agenda on industry needs to take account of the other dimensions of the industry policy consolidation; and the economic debate needs to deepen its study of cohesion between industrial and macroeconomic policies, following a strategy for long-term national development.

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Annex A1
Classification of the sectors by type of technology, according to OECD

The sectors are classified by type of technology based on the classification published by OECD (2005). As some of those sectors form part of the group of 56 categories used by the IBGE System of National Accounts since 2000, the weight of production of each sector in the group of 56 categories was calculated for the corresponding sector in the group of 42. Those with the heavier weights determined how the sectors were classified in the latter group.

For example, the sectors "Wood products, excluding furniture" and "Furniture and products of miscellaneous industries" of the group of 56 categories correspond to "Wood and furniture and miscellaneous industries" in the group of 42. The sector "Wood products, excluding furniture" is natural-resource-intensive, whereas "Furniture and products of miscellaneous industries" is labour-intensive. As the weight of the sector "Furniture and products of miscellaneous industries" was greater (69.71%), "Wood and furniture and miscellaneous industries" were classified as labour-intensive.

Table A1.1
Classification of the sectors by type of technology, according to OECD

Group of 42 categories	Group of 56 categories	Weight of production of the sector of the group of 56 categories in the group of 42	OECD classification of the sector of the group of 56 categories	OECD classification of the sector of the group of 42 categories
Pharmaceutical and veterinary	Perfumery hygiene and cleaning	40.57%	Scale-intensive	
	Pharmaceutical products	59.43%	Science-based	Science-based
Extractive mineral	Iron ore	22.60%	Natural-resource-intensive	Natural-resource-intensive
	Other products of the extractive industry	14.92%	Natural-resource-intensive	
Oil and gas	Oil and natural gas	62.48%	Natural-resource-intensive	Natural-resource-intensive
Non-metallic mineral	Other non-metallic mineral products	77.25%	Natural-resource-intensive	Natural-resource-intensive
	Cement	22.75%	Natural-resource-intensive	
Oil refining	Oil refining and coke	58.05%	Natural-resource-intensive	Natural-resource-intensive
	Alcohol	8.68%	Natural-resource-intensive	
Chemical elements	Chemicals	24.93%	Scale-intensive	Scale-intensive
	Resin and elastomer manufacture	8.34%	Scale-intensive	
Coffee industry	Food and beverages	96.92%	Natural-resource-intensive	Natural-resource-intensive
Processing of vegetable products	Tobacco products	3.08%	Natural-resource-intensive	Natural-resource-intensive
Animal slaughtering				Natural-resource-intensive
Dairy products industry				Natural-resource-intensive
Sugar manufacture				Natural-resource-intensive

Table A1.1 (concluded)

Group of 42 categories	Group of 56 categories	Weight of production of the sector of the group of 56 categories in the group of 42	OECD classification of the sector of the group of 56 categories	OECD classification of the sector of the group of 42 categories
Vegetable oil manufacture				Natural-resource-intensive
Other food products				Natural-resource-intensive
Wood and furniture	Wood products, excluding furniture	30.29%	Natural-resource-intensive	Labour-intensive
Miscellaneous industries	Furniture and products of miscellaneous industries	69.71%	Labour-intensive	Labour-intensive
Textile industry	Textiles	100.00%	Labour-intensive	Labour-intensive
Garments	Garments and clothing accessories	100.00%	Labour-intensive	Labour-intensive
Footwear manufacture	Leather articles and footwear	100.00%	Labour-intensive	Labour-intensive
Iron and steel	Manufacture of steel and derivatives	100.00%	Scale-intensive	Scale-intensive
Other metallurgical Non-ferrous metallurgical	Metal products except machinery and equipment	67.30%	Labour-intensive	Labour-intensive
	Non-ferrous metallurgy	32.70%	Natural-resource-intensive	Natural-resource-intensive
Automobiles, trucks, buses	Automobiles, trucks and utility vehicles	79.96%	Scale-intensive	Scale-intensive
	Trucks and buses	20.04%	Scale-intensive	
Vehicle parts	Automobile parts and accessories	66.12%	Scale-intensive	Scale-intensive
	Other transport equipment	33.88%	Scale-intensive	
Rubber industry	Articles of rubber and plastic	100.00%	Scale-intensive	Scale-intensive
Plastics				
Miscellaneous chemicals	Miscellaneous chemical products and preparations	33.70%	Scale-intensive	Scale-intensive
	Pesticides	38.14%	Scale-intensive	
	Paints, varnishes, enamels and lacquer	28.16%	Scale-intensive	
Machinery and equipment	Machinery and equipment, including maintenance and repair	100.00%	Differentiated	Differentiated
Electrical material	Household electrical appliances	11.94%	Differentiated	Differentiated
Electronic equipment	Office machinery and information technology equipment	16.70%	Science-based	Differentiated
	Electrical machines, apparatus and materials	35.92%	Differentiated	
	Electronic materials and communication equipment	23.16%	Differentiated	
	Medical-hospital apparatus and instruments, of measurement and optics	12.28%	Differentiated	

Source: Prepared by the authors, on the basis of Organization for Economic Cooperation and Development (OECD), *Science, Technology and Industry Scoreboard 2005*, Paris, 2005; and K. Pavitt, "Sectoral patterns of technical change: towards a taxonomy and a theory", *Research Policy*, vol. 13, No. 6, Amsterdam, Elsevier [online] http://www.sourceoecd.org/scienceIT/9264010556.

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