The effects of domestic demand shocks on inflation in a small open economy: Chile in the period 2000–2021

Ramón Eugenio López and Kevin Alejandro Sepúlveda

Abstract

This study decomposes the factors that determined inflation in Chile during the period 2000–2021. It finds that the main determinants of domestic inflation were variables of external origin and the exchange rate. Domestic demand played a rather limited role as an inflationary factor. In normal periods, increases in domestic demand generally explained no more than 25% of observed inflation. The average monthly inflation observed during the period 2000–2021 was 0.3%, which means that domestic demand growth in normal periods explained monthly inflation of 0.08%. Surprisingly, the extraordinary periods of rapid demand growth resulting from highly expansionary fiscal policies, large withdrawals from pension retirement savings or both had a rather modest effect on inflation. This study corroborates what might be expected in a small, open economy like Chile's: most domestic price changes are determined by foreign price changes.

Keywords

Inflation, supply and demand, prices, foreign exchange rates, liquidity, monetary policy, econometric models, Chile

JEL classification

E31, E5, E62

Authors

Ramón Eugenio López holds a PhD in Economics and is a professor at the Department of Economics of the University of Chile. Email: ramlopez@fen.uchile.cl.

Kevin Alejandro Sepúlveda holds a Master of Science in Economics and is a co-researcher at the Department of Economics of the University of Chile. Email: kesepulved@fen.uchile.cl.

I. Introduction

This paper studies the determinants of inflation in Chile during the first two decades of the twenty-first century. Its main objective is to assess the importance of fluctuations in domestic demand for domestic inflation. This is a very significant issue, considering that Keynesian policies in a context of recession, low levels of production capacity utilization and high unemployment boost domestic spending through higher public expenditure, higher social subsidies and related measures such as reductions in the compulsory saving requirements imposed on workers under the system of pension fund administrators (AFPs). Precisely these policies were implemented very aggressively in Chile during the coronavirus disease (COVID-19) crisis over the period 2020–2021.

Determining the inflationary effect of such Keynesian policies is of vital importance for the future design of demand policies. The idea behind Keynesianism is that, under conditions of production capacity underutilization and unemployment, increases in demand translate into greater use of and even expansion of production capacity rather than into higher prices. If the effect of expanding demand is fundamentally to generate higher inflation, we can conclude that demand policies are of limited value. This is what motivates this research. In 2020 and 2021, policies designed to expand demand at rates rarely seen in the recent history of Chile were implemented in the country. Our aim is to take advantage of this gigantic experiment to evaluate the relevance of demand policies in an economy like Chile's under crisis conditions.[1]

This paper adopts an eclectic position which allows for the possibility that domestic demand factors may affect inflation in confluence with external inflation. The determinants of changes in domestic inflation are specified as changes in inflation abroad, movements in the nominal exchange rate and internal demand factors. Since the exchange rate is likely to be affected by demand factors, we provide for simultaneous determination of inflation and the exchange rate.

II. The model

We start with a basic price equation which we subsequently modify to deal with several potential econometric issues and to include some additional dynamic considerations:

$$p_t = AP_t^{\alpha} E_t^{\beta} D_t^{\gamma} e^{\mu_t} \tag{1}$$

where p_t is the level of the domestic price index in month t, A is a constant, P_t is the foreign price index, E_t is the nominal exchange rate, D_t is the level of internal demand, α,β,γ are parameters and μ_t is the statistical error.

We first express equation (1) in logarithms. Additionally, to account for the possibility of unit roots, we estimate this equation using log first differences:

$$\Delta lnp_t = \alpha\Delta lnP_t + \beta\Delta lnE_t + \gamma\Delta lnD_t + \Delta\mu_t \tag{2}$$

[1] Chile is a small economy that is extremely open to both trade and capital flows. This would suggest in principle that if the purchasing power parity (PPP) hypothesis holds even in its weakest versions, such as relative parity, or even if a quasi-PPP condition prevails (Hegwood and Papell, 1998), the inflationary effect of demand policies, such as those implemented in 2020 and 2021, should be limited. This in turn would imply that domestic demand stimuli should be effective in promoting production and employment and possibly investment, with only a small inflationary effect. In other words, well-implemented demand policies in a country affected by a chronic lack of demand caused mainly by high income concentration and low wages could trigger a spiral of growth and development over the medium term. The evidence regarding the empirical validity of the PPP hypothesis in any of its versions is mixed. See studies by Taylor (2002), Crownover, Pippenger and Steigerwald (1996), Li, Lin and Hsiao (2015), Carlsson, Lyhagen and Österholm (2007), Kasuya and Ueda (2000), Céspedes and De Gregorio (1999), Korap and Aslan (2010) and Taylor and McMahon (1988).

$$\text{where } \Delta lnx_t \equiv lnx_t - lnx_{t-1} \text{ for } x_t = \begin{cases} P_t \\ E_t \\ D_t \end{cases} \tag{3}$$

We use as proxies for the variable ΔlnP_t, inflation in Chile's most important trading partners, i.e., the United States, the European Union and China. Changes in the exchange rate, ΔlnE_t, are measured monthly in Chile. Since the variable ΔlnD_t is more difficult to measure, we assume that in normal periods domestic demand grows at a stable rate consistent with the long-term growth of the economy, while in periods of crisis domestic demand may be massively affected by exceptional fiscal policies such as unusual increases or reductions in social subsidies or other policies such as the withdrawal of savings held by the population in the AFP pension funds or in unemployment funds, among others. Thus, for the purposes of econometric estimation, equation (2) is adjusted as follows:

$$\Delta lnp_t = \gamma_0 + \sum_1^n \gamma_i d_i + \alpha \Delta lnP_t + \beta \Delta lnE_t + \Delta\mu_t \tag{4}$$

where d_i $(i = 1, 2, \ldots n)$ are dummy variables with values equal to 1 in period i and equal to zero in periods $j \neq i$. The effect of demand changes on inflation, $\gamma \Delta lnD_t$, can be decomposed into an effect in periods of stability, when demand increases at a stable rate γ_0, and an effect in abnormal or critical periods, when it is influenced by emergency fiscal or monetary measures to deal with particular crises, with γ_i $(i = 1, 2 \ldots n)$, where n abnormal periods are considered. Therefore, the effect of changes in demand on inflation in period i is $\gamma \Delta lnD_t = \gamma_0 + \sum_1^n \gamma_i d_i$. We also consider an alternative specification for the demand factors, using a continuous proxy indicator of monthly demand which assumes it to change proportionally with the Monthly Index of Economic Activity (IMACEC). In other words, we assume that ΔlnD_t is positively correlated with $\Delta IMACEC$ (the result using this specification is reported in annex A1).

Equation (4) is also augmented with past values of the dependent variable on the right-hand side to capture possible lagged effects of the variables on inflation. This makes it possible to obtain measures of the short- and medium-term effects of the independent variables on the level of inflation.

1. Estimation methods

The estimation of equation (4) may be affected by simultaneity bias. Two alternative methods of estimating the expanded equation (4) are used. One is two-stage least squares (TSLS), employing instrumental variables for the exchange rate to avoid simultaneity biases arising from the fact that this can be affected by domestic inflation. We use the price of copper and the United States Federal Funds Effective Rate (FFER) as instrumental variables. The generalized method of moments (GMM) is used in addition to the TSLS estimates and can be regarded as a test of the robustness of the TSLS-estimated coefficients.

2. Data

These equations are estimated using monthly data for the period 2000–2021, during which there were no major structural changes affecting the determination of the exchange rate, so that the Taylor (1988) requirement for the validity of PPP in any of its forms is met. The data used are those provided by official institutions. In particular, the inflation data for Chile, the United States, China, and the eurozone were obtained, respectively, from the National Institute of Statistics (INE), the Bureau of Labor Statistics (BLS), the National Bureau of Statistics (NBS) and the European Statistical Office (Eurostat). The data on the nominal exchange rate, the monetary policy rate (MPR) and the Monthly Index of Economic Activity (IMACEC) were obtained from the Statistical Database of the Central Bank of Chile. Lastly, the copper price was taken from the Chilean Copper Commission (COCHILCO) and the United States Federal Funds Effective Rate (FFER) from the Federal Reserve. Annex A1 presents complete descriptive statistics for the data used in the regressions.

3. Natural experiments and dummy variables

Model (4) includes six dummy variables for abnormal periods. The first (d_1) (takes the value 1 for the period between September 2008 and September 2009. This dummy is used to capture the effect of the international crisis on Chile in those months, when there were major changes in both production and domestic monetary and fiscal policies. The second dummy variable (d_2) takes the value 1 for the months between October 2009 and February 2010, the period of recovery from the crisis.

The third dummy variable (d_3) takes the value 1 for the period between April 2020 and July 2020, a time when the economic effects of the pandemic were present but significant social subsidies and AFP withdrawals had yet to be implemented. The fourth dummy variable (d_4) takes the value 1 for the period between August 2020 and December of the same year, a period in which the effects of the pandemic coincided with those of the first social subsidies and the first AFP withdrawal in July 2020. The fifth dummy variable (d_5) takes the value 1 for the period between January 2021 and April 2021, when social subsidies were greatly increased and the second AFP withdrawal in December 2020 had its effect on domestic demand. Lastly, the sixth dummy variable (d_6) takes the value 1 for the period between May 2021 and December 2021, when the third AFP withdrawal in April 2021 and large social subsidies such as the Emergency Family Income (IFE) produced their effects.[2]

Thus, these dummies capture the impact of the various natural experiments that took place in 2020 and 2021. They show how far-reaching the impact of these was, including that of the pandemic between April and July 2020 when no policy protection was as yet in place (d_3). Then there is a period when the experiment includes the combined impact of the pandemic and the earliest, very limited social subsidies (d_4). The clearest demand effects ought to be captured by the coefficients of d_5 and especially d_6, these being the periods when the unusual demand stimuli operated very intensively and persistently. Indeed, the stimulus to demand at this time was the greatest in the country's recent history. If demand stimuli cause higher inflation, then these coefficients should be expected to be positive and significant.

III. Results

This section presents the results of the TSLS and GMM estimates. Table 1 shows the results of the second stage of the estimation by instrumental variables, where the variable explained is monthly inflation.

Table 1
Monthly inflation estimated
by two-stage least squares

Number of observations	263
Wald chi-squared (11)	209.27
Probability > chi-squared	0.0000
R-squared	0.4357

Chile inflation	Coefficient	Robust standard error	z	$p > \|z\|$	Confidence interval (95%)	
Percentage change in nominal exchange rate	0.0601	0.0133	4.53	0	0.0341	0.0861
United States inflation	0.4481	0.0528	8.49	0	0.3447	0.5515
China inflation	0.0306	0.0296	1.03	0.302	-0.0275	0.0886
Eurozone inflation	0.0371	0.0428	0.87	0.386	-0.0467	0.1209

2 This was also the period when the central bank monetary policy rate (MPR) started to rise quite rapidly. In annex A1, we explicitly use changes in the MPR with lags as an additional explanatory variable.

Table 1 (concluded)

Chile inflation	Coefficient	Robust standard error	z	$p > \|z\|$	Confidence interval (95%)	
Dummy 1	-0.0006	0.0014	-0.42	0.675	-0.0032	0.0021
Dummy 2	-0.0014	0.0014	-1.01	0.315	-0.0042	0.0013
Dummy 3	-0.0003	0.0007	-0.51	0.613	-0.0016	0.0010
Dummy 4	0.0017	0.0013	1.27	0.204	-0.0009	0.0043
Dummy 5	-0.0004	0.0014	-0.28	0.779	-0.0032	0.0024
Dummy 6	-0.0002	0.0016	-0.15	0.882	-0.0033	0.0028
Chile inflation with a one-period lag	0.3694	0.0566	6.52	0	0.2584	0.4804
γ_0	0.0007	0.0003	2.67	0.008	0.0002	0.0012

Source: Prepared by the authors.

Important aspects of the results in table 1 are the following. First, the impact of external inflation, especially that of the United States, is very large and significant. Second, the exchange rate is another variable that plays a quantitatively important and significant role. Third, the low significance of the demand variables is striking, the exception being the coefficient of the constant, which can be interpreted as the effect of movements in demand in normal periods, and which is statistically significant at 95%. The dummy variables that capture possible inflationary changes due to an acceleration of demand growth in particular periods are not statistically significant.

One interpretation of the demand estimators in table 1 is that increases in demand in normal periods explained an average of 0.07% of the monthly inflation rate when the average monthly rate in normal periods was 0.29%, i.e., demand growth explained 24% of inflation in normal periods. Thus, although normal demand growth does seem to have some effect on inflation, the effect of demand in exceptional periods, as captured by the six dummy variables, does not seem to have played any additional role in inflation.

However, because the exchange rate is kept constant, the estimates in table 1 show only the direct effect of these variables on inflation. It is possible that a significant part of the inflationary effect of some of these variables and of demand occurs through their impact on the exchange rate. We capture these indirect effects in the first stage of the TSLS estimation that we present below. Table 2 shows the estimators of the first stage of the TSLS regression, where the exchange rate is the variable that is explained by the exogenous variables plus the instruments.

Table 2

First stage estimators: the exchange rate

Number of observations	263
F (12, 250)	6.16
Probability > F	0.0000
R-squared	0.3242
Adjusted R-squared	0.2917

Percentage change in nominal exchange rate	Coefficient	Robust standard error	t	$p > \|t\|$	Confidence interval (95%)	
United States inflation	0.0949	0.4508	0.21	0.833	-0.7929	0.9828
China inflation	-0.3179	0.2362	-1.35	0.179	-0.7831	0.1472
Eurozone inflation	0.0857	0.3319	0.26	0.796	-0.5679	0.7394
Dummy 1	-0.0008	0.0100	-0.08	0.936	-0.0206	0.0190
Dummy 2	-0.0032	0.0167	-0.19	0.847	-0.0360	0.0296
Dummy 3	-0.0115	0.0088	-1.31	0.192	-0.0289	0.0058

Table 2 (concluded)

Percentage change in nominal exchange rate	Coefficient	Robust standard error	t	$p > \|t\|$	Confidence interval (95%)	
Dummy 4	-0.0075	0.0074	-1.01	0.313	-0.0221	0.0071
Dummy 5	-0.0048	0.0078	-0.62	0.535	-0.0202	0.0105
Dummy 6	0.0193	0.0068	2.84	0.005	0.0059	0.0327
Chile inflation with a one-period lag	0.1096	0.4441	0.25	0.805	-0.7650	0.9842
Percentage change in copper price	-0.2107	0.0379	-5.56	0	-0.2853	-0.1361
Percentage change in Federal Funds Effective Rate	-0.0160	0.0082	-1.95	0.052	-0.0321	0.0001
γ_0	0.0038	0.0020	1.96	0.051	0.0000	0.0077

Source: Prepared by the authors.

As can be seen in table 2, only dummy 6 has a positive and statistically significant effect on the exchange rate. This dummy represents the period from May 2021 to December of the same year during which demand accelerated because of the combination of fiscal demand stimuli and the third AFP withdrawal.[3] The coefficient of this dummy suggests that the expansion of demand explained a 1.93% depreciation in the monthly nominal exchange rate, a large effect that indirectly impacted inflation, increasing it by a net 0.12% per month. In the second half of 2021, in other words, the demand effect explained almost 20% of the average inflation in that period, which was 0.68% per month.

In addition to the TSLS estimation, we use the GMM method, whose results can be treated as part of the robustness analysis of the TSLS estimators. Table 3 provides the results obtained with the GMM method.

Table 3
Monthly inflation estimated by the generalized method of moments

Number of observations	263
Wald chi-squared (11)	211.05
Probability > chi-squared	0.0000
R-squared	0.4357

Chile inflation	Coefficient	Robust standard error	z	$p > \|z\|$	Confidence interval (95%)	
Percentage change in nominal exchange rate	0.0601	0.0132	4.53	0	0.0341	0.0860
United States inflation	0.4480	0.0527	8.51	0	0.3448	0.5512
China inflation	0.0306	0.0296	1.03	0.302	-0.0275	0.0886
Eurozone inflation	0.0371	0.0428	0.87	0.386	-0.0467	0.1209
Dummy 1	-0.0006	0.0014	-0.42	0.675	-0.0032	0.0021
Dummy 2	-0.0014	0.0014	-1.01	0.314	-0.0041	0.0013
Dummy 3	-0.0003	0.0007	-0.51	0.609	-0.0016	0.0009
Dummy 4	0.0017	0.0013	1.27	0.204	-0.0009	0.0043
Dummy 5	-0.0004	0.0014	-0.28	0.779	-0.0032	0.0024
Dummy 6	-0.0002	0.0016	-0.15	0.883	-0.0033	0.0028
Chile inflation with a one-period lag	0.3694	0.0566	6.53	0	0.2585	0.4803
γ_0	0.0007	0.0003	2.68	0.007	0.0002	0.0012

Source: Prepared by the authors.

[3] There were also large adjustments in the monetary policy rate (MPR) in this period. We report this same estimate in annex A1, but separating out the effect of the MPR.

As can be seen by comparing tables 1 and 3, the values and significance of the parameters estimated by TSLS and GMM are quite similar. In both estimates, the most important variables affecting domestic inflation are external inflation, particularly that of the United States, and the exchange rate. On the other hand, the direct effect of domestic demand seems to be of little significance. However, given the importance of the exchange rate as a determinant of inflation, as stated above, it is possible that the effect of demand on inflation occurs fundamentally through this variable. For this reason, we focus on estimating the first stage of TSLS in order to include the possible indirect effect of demand on the exchange rate as a factor driving inflation.

1. The direct effect of domestic demand

Using the estimated parameters, we find that the estimated effect of domestic demand is γ_0 in a normal period, while in the abnormal period i it is given by $\gamma_0 + \gamma_i$, where $i = 1, 2..., 6$ is associated with the period in which the dummy d_i takes the value 1.

The direct effect of the demand stimuli considered is very small and of low statistical significance in all the periods. The figures presented in tables 1 to 3 consider only the individual statistical significance of each variable. In addition, we consider the joint significance of all demand variables on inflation.

2. Joint significance tests on demand effects

We perform a joint significance test on the constant and the parameters of the dummy variables identified for the years 2020 and 2021. Specifically, the null hypothesis for both estimation methods is:

$$H_0 : \gamma_0 = \gamma_3 = \gamma_4 = \gamma_5 = \gamma_6 = 0 \tag{5}$$

The tests yielded p-values of 0.0588 and 0.0585 for the TSLS and GMM estimations, respectively. This implies that we can reject with 90% confidence the null hypothesis that the parameters associated with these dummy variables and the constant are simultaneously zero. However, we cannot reject this hypothesis with 95% confidence. Thus, the direct effect of demand factors on inflation generally appears to be of weak statistical significance. We also perform the following joint significance tests:

$$H_0 : \gamma_0 = \gamma_i = 0, i = 1,...,6 \tag{6}$$

With the results of the TSLS and GMM estimations, we can reject this hypothesis for $i = 1,...,6$ with 95% confidence.

3. Total effect of demand on inflation

The total effect of demand is calculated using the parameters estimated by TSLS. It consists of the partial effect of demand on inflation at a given exchange rate, obtained directly from the demand parameters estimated in the second stage of the TSLS estimation (table 1), plus the indirect effect of demand on the exchange rate, which is obtained using the parameters estimated in the first stage of the TSLS estimation. Thus, the total effect of demand on inflation in period i is

$$\Delta Inf_i = \gamma_0 + \gamma_i + \beta(\varepsilon_0 + \varepsilon_i) \tag{7}$$

where ΔInf_i is the total effect on inflation caused by the increase in demand in period i, β is the second stage parameter that measures the impact of the exchange rate on inflation, and $\varepsilon_0 + \varepsilon_i$ is the effect of demand on the exchange rate in period i, obtained by estimating the equation of the first stage shown in table 2. We use the TSLS estimators to calculate the standard error of equation (7).

A disadvantage of this method, though, is that it imposes values of zero for the covariances between the parameters estimated in the separate equations. For this reason, we also estimate the model using three-stage least squares (not shown in the tables), which does allow us to derive the covariances between different equations.

Thus, the impact of demand on inflation is augmented when exchange-rate effects are considered, but these effects are generally not significant, as can be seen in table 4. Only for the periods from August to December 2020 and May to December 2021 is the p-value close to having some statistical significance. Using equation (7), the total average monthly effect of demand is found to be close to 0.22% for the five-month period from August to December 2020, when the inflationary effect of demand was highest, and around 0.19% for the period from May to December 2021. Given that monthly inflation in the period from May to December 2021 averaged 0.68%, the impact of domestic demand explains around 28% of observed inflation in that period. Thus, if monetary policy had completely suppressed the expansion of domestic demand, total inflation in 2021 would have been 5.2% instead of 7.2%.

Table 4
Total effect of domestic demand on inflation and p-values under the null hypothesis that the respective coefficients are equal to zero

	$i = 1$	$i = 2$	$i = 3$	$i = 4$	$i = 5$	$i = 6$
Coefficient	0.000307	-0.000674	-0.000100	0.002168	0.000231	0.001856
p-value (two-stage least squares)	0.7344	0.6463	0.9514	0.1409	0.8911	0.1364
p-value (three-stage least squares)	0.7150	0.6214	0.9478	0.1131	0.8828	0.1090

Source: Prepared by the authors.

IV. Simulations

The parameters estimated by TSLS were used to simulate average inflation for the whole of 2021 and the last quarter of the same year. To this end, data on external inflation in the United States, China and the eurozone, lagged inflation in Chile and the change in the nominal exchange rate were included.

As can be seen in table 5, the model is capable of simulating the inflationary events of 2021 fairly accurately. This provides some reassurance about the explanatory capacity of the model. Annex A1 provides additional simulations for each month in 2021 and also for the months of January and February 2022, which are not part of the sample used to estimate the model.

Table 5
Contributions to the average inflation rate simulated using parameters estimated by two-stage least squares, whole of 2021 and last quarter of 2021
(Percentages)

Average inflation in fourth quarter of 2021	Percentage change in nominal exchange rate	United States inflation	China inflation	Eurozone inflation	Chile inflation with a one-period lag	Simulation: predicted value
0.867	0.164	0.243	0.008	0.020	0.369	0.805
Average inflation in 2021	**Percentage change in nominal exchange rate**	**United States inflation**	**China inflation**	**Eurozone inflation**	**Chile inflation with a one-period lag**	**Simulation: predicted value**
0.592	0.074	0.255	0.004	0.015	0.203	0.551

Source: Prepared by the authors.

V. The vector autoregressive model

We use a vector autoregressive (VAR) model to characterize the simultaneous dynamic interactions between domestic inflation, changes in the nominal exchange rate and changes in the monetary policy rate (MPR). We start from a general model of the form

$$y_t = \nu + \sum_{i=1}^{p} A_i y_{t-i} + B x_t + \mu_t \tag{8}$$

where $y_t = (y_{1t,\dots} y_{kt})'$ is a $(k \text{ x } 1)$ vector containing the changes over time in the k endogenous variables; A_i are $(k \text{ x } k)$ arrays of parameters; x_t is a $(k_e \text{ x } 1)$ vector of exogenous variables; B is a $(k \text{ x } k_e)$ matrix of coefficients; ν is a $(k \text{ x } 1)$ parameter vector; and μ_t is assumed to be white noise, i.e., $E(_{\mu t}) = 0$; $E(\mu_t \mu_t') = \Sigma$ and $E(\mu_t \mu_s') = 0$ for $t \neq s$.

To select the lag model that best fits the data, we applied the Bayesian information criterion (BIC). There is a variety of evidence that this criterion is among those with the best predictive performance (Lütkepohl, 1985; Clark, 2004). We also used the Akaike information criterion (AIC) and found, after applying different specifications for model (8), that both criteria suggested that the model best fitting the data was a VAR (1). Accordingly, we selected the following model for the estimates:[4]

$$y_t = \nu + A_1 y_{t-1} + Bx_t + \mu_t \tag{9}$$

where y_t is a $(3 \text{ x } 1)$ vector of the endogenous variables considered, i.e., domestic inflation, the change in the nominal exchange rate and the change in the MPR. In addition, the vector x_t contains the inflation of the United States, China and the eurozone. Analysis of the stability of specification (9) showed that all the eigenvalues were within the unit circle, meaning that the model was stable and we could proceed to carry out analyses on the results obtained in the estimations.

The next subsection presents the results of the estimates and the dynamic relationships between domestic inflation, changes in the nominal exchange rate and changes in the MPR, analysing the orthogonalized impulse-response functions (IRFs) of a shock in the MPR and a shock in the nominal exchange rate as they affect inflation. As Nguyen, Papyrakis and Van Bergeijk (2019) point out, IRFs predict the sign, magnitude and statistical significance of responses to shocks in the variables of interest.[5] These orthogonalized IRFs are calculated using the Cholesky decomposition, which allows us to analyse the direct effect of the shock on inflation. Additionally, Granger causality tests are performed and the prediction error is decomposed.

1. Data

The data used in the estimates are the same as were used in the model in the previous section, i.e., monthly first difference data for the period 2000–2021.

2. Vector autoregression results

The results of the model (9) estimations are presented in table 6.

[4] It is important to note that the main results presented in the next subsection are not qualitatively altered when the other specifications evaluated with these criteria are considered.

[5] These authors develop a VAR model using monthly data to measure the effects of monetary policy on the Vietnamese economy.

Table 6
First-order vector autoregression estimates

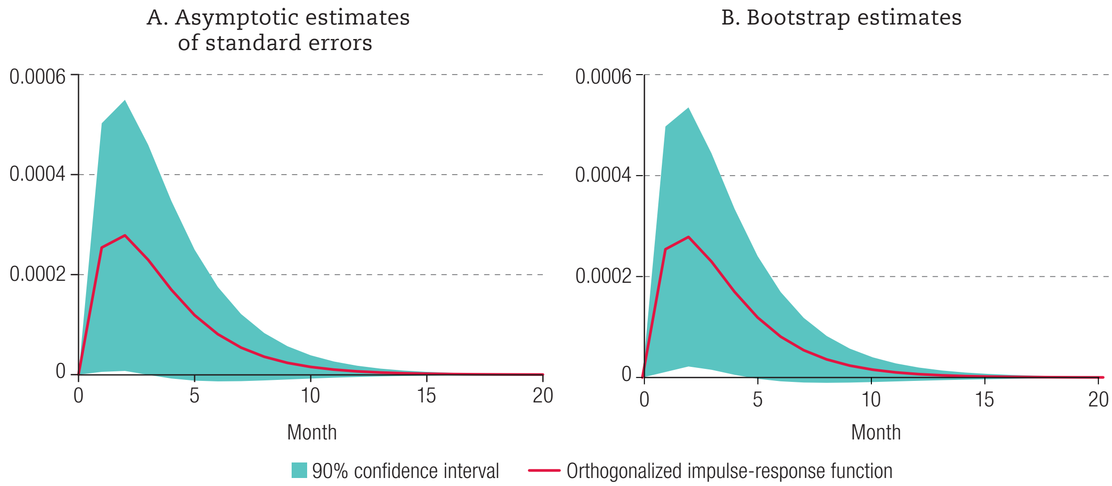
	Coefficient	Standard error	z	$p > \|z\|$	Confidence interval (90%)	
Percentage change in nominal exchange rate (NER)						
Percentage change in NER with a one-period lag	0.2670208	0.064315	4.15	0	0.161233	0.372809
Chile inflation with a one-period lag	0.37395	0.44234	0.85	0.398	-0.35363	1.101535
Percentage change in monetary policy rate (MPR) with a one-period lag	0.0093916	0.012409	0.76	0.449	-0.01102	0.029803
United States inflation	-0.6543436	0.454322	-1.44	0.150	-1.40164	0.092949
China inflation	-0.3334025	0.264478	-1.26	0.207	-0.76843	0.101625
Eurozone inflation	0.1397398	0.393051	0.36	0.722	-0.50677	0.78625
v_1	0.0022626	0.00216	1.05	0.295	-0.00129	0.005816
Chile inflation						
Percentage change in NER with a one-period lag	0.0280341	0.007919	3.54	0	0.015009	0.041059
Chile inflation with a one-period lag	0.3557654	0.054464	6.53	0	0.26618	0.445351
Percentage change in MPR with a one-period lag	0.0025765	0.001528	1.69	0.092	6.33E-05	0.00509
United States inflation	0.4315452	0.055939	7.71	0	0.339533	0.523557
China inflation	0.0098239	0.032564	0.30	0.763	-0.04374	0.063388
Eurozone inflation	0.0450226	0.048395	0.93	0.352	-0.03458	0.124626
v_2	0.0008119	0.000266	3.05	0.002	0.000374	0.001249
Percentage change in MPR						
Percentage change in NER with a one-period lag	0.5670869	0.262242	2.16	0.031	0.135737	0.998437
Chile inflation with a one-period lag	-0.6430093	1.80364	-0.36	0.721	-3.60973	2.323715
Percentage change in MPR with a one-period lag	0.6380204	0.050599	12.61	0	0.554792	0.721249
United States inflation	2.480006	1.852495	1.34	0.181	-0.56708	5.52709
China inflation	0.2028631	1.078407	0.19	0.851	-1.57096	1.976684
Eurozone inflation	-0.8832105	1.602662	-0.55	0.582	-3.51935	1.752933
v_3	-0.0000129	0.008809	0.00	0.999	-0.0145	0.014477

Source: Prepared by the authors.
Note: v_i $(i = 1,2,3)$ are constants.

We shall now analyse the effect of shocks in the MPR and exchange rate on inflation.

Figure 1 shows the effect on inflation when the change in the MPR unexpectedly increases by one standard deviation. The graphs show that this shock has a small but positive and statistically significant effect on inflation in the short term, peaking in about the third month, after which the effect tails off and ceases to be statistically significant. Thus, an unexpectedly large rise in the MPR is associated with a small but perceptible increase in inflation in the short term. This is surprising, since one would expect a rise in the MPR to have a negative, not a positive, effect on inflation. One possible explanation is that in a small, open economy with inflation generated abroad, the effect of a rise in the MPR translates into an increase in the cost of imported products due to the higher cost of inventories and of imported intermediate goods, and the effect of this increase in pushing up inflation may be stronger than the effect of the rise in the MPR in reducing domestic demand.

Figure 1
Orthogonalized impulse-response function for the effects of a monetary policy rate shock of one standard deviation on inflation



Source: Prepared by the authors.

Another possible explanation is that, since the MPR is a variable controlled by the central bank and increases in this variable are due to its expectations about future inflation, such an increase may reflect a short-term rise in inflation that was going to occur anyway if, on average, these expectations are correct.

Since the estimates arrived at using the Cholesky decomposition are not necessarily invariant to the order of the endogenous variables, figure 2 presents the IRFs for different orders. In this graph, we can observe that the qualitative results shown in figure 1 change only for one order (in which the effect is not statistically significant), while for some other orders of the Cholesky decomposition the increase in inflation is instantaneous and falls monotonically to zero as the months pass, by contrast with the pattern in figure 1.

Figure 2
Orthogonalized impulse-response function for the effects of a monetary policy rate shock of one standard deviation on inflation for different orders in the Cholesky decomposition

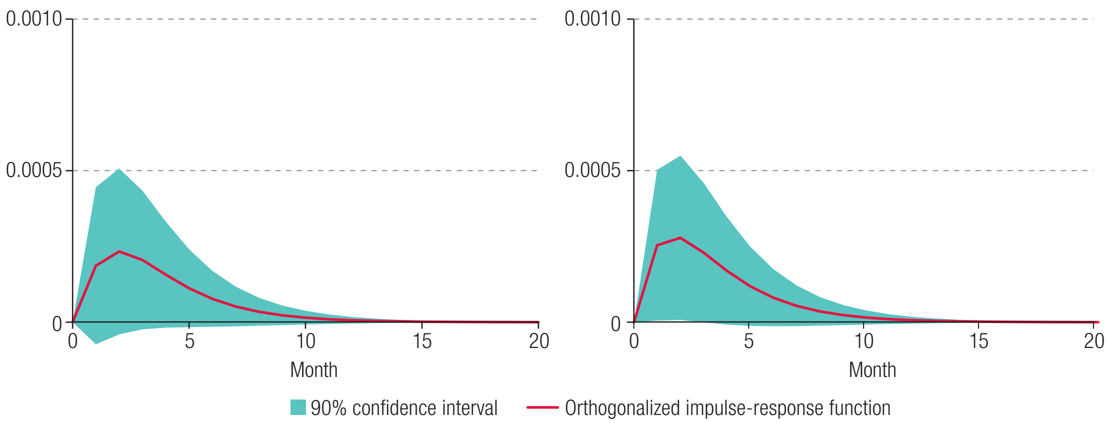
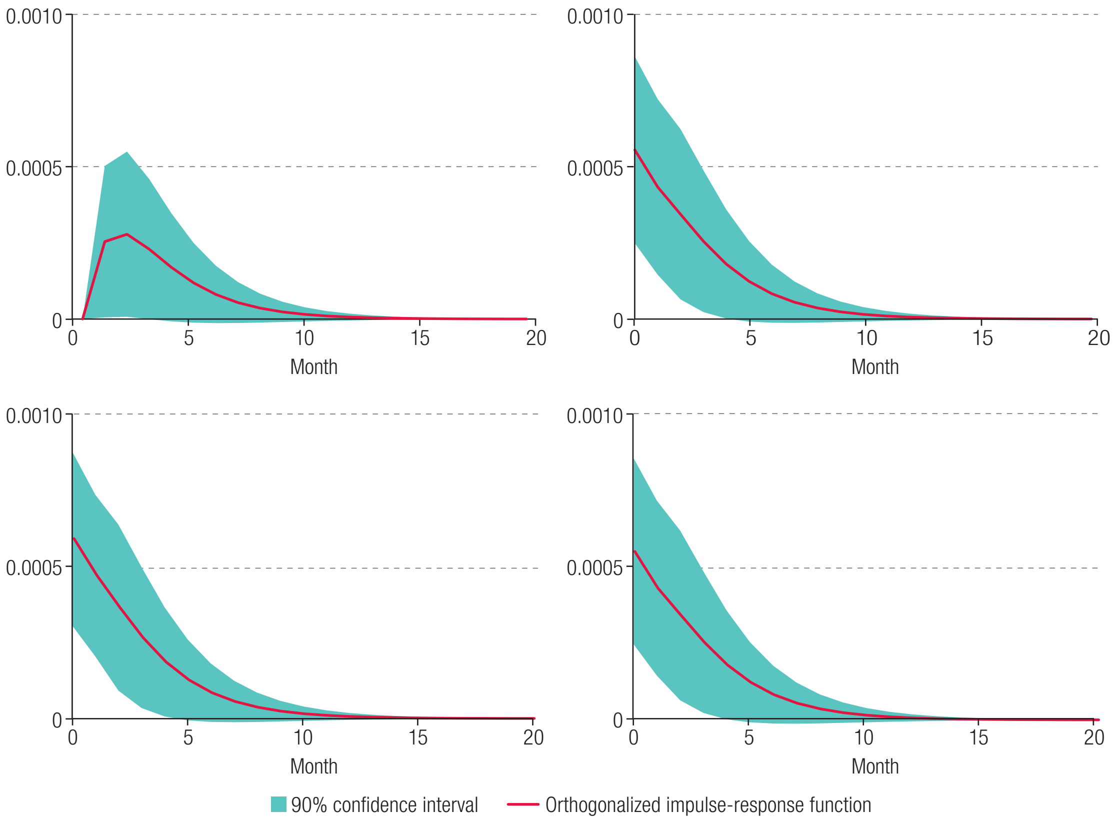


Figure 2 (concluded)



Source: Prepared by the authors.

The effect of an unexpectedly large movement in the exchange rate also depends on the order of the Cholesky decomposition. For some orders, as figure 3 shows, the effect on inflation arises instantaneously and increases before falling back to zero as the months go by, while for other orders there is a rising effect that peaks in the second month and then falls to zero in about the seventh month. However, the effect is always positive and statistically significant.

Figure 3
Orthogonalized impulse-response function for the effects of an exchange-rate shock of one standard deviation on inflation for different orders in the Cholesky decomposition

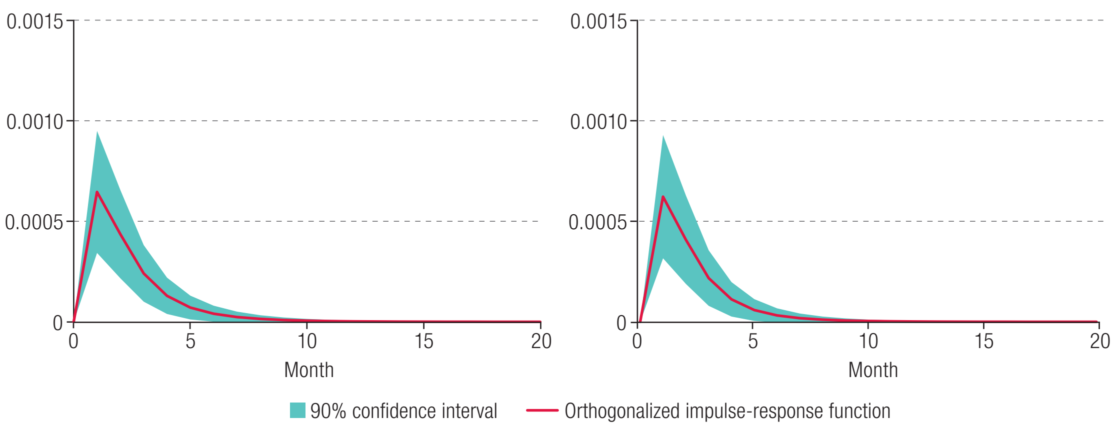
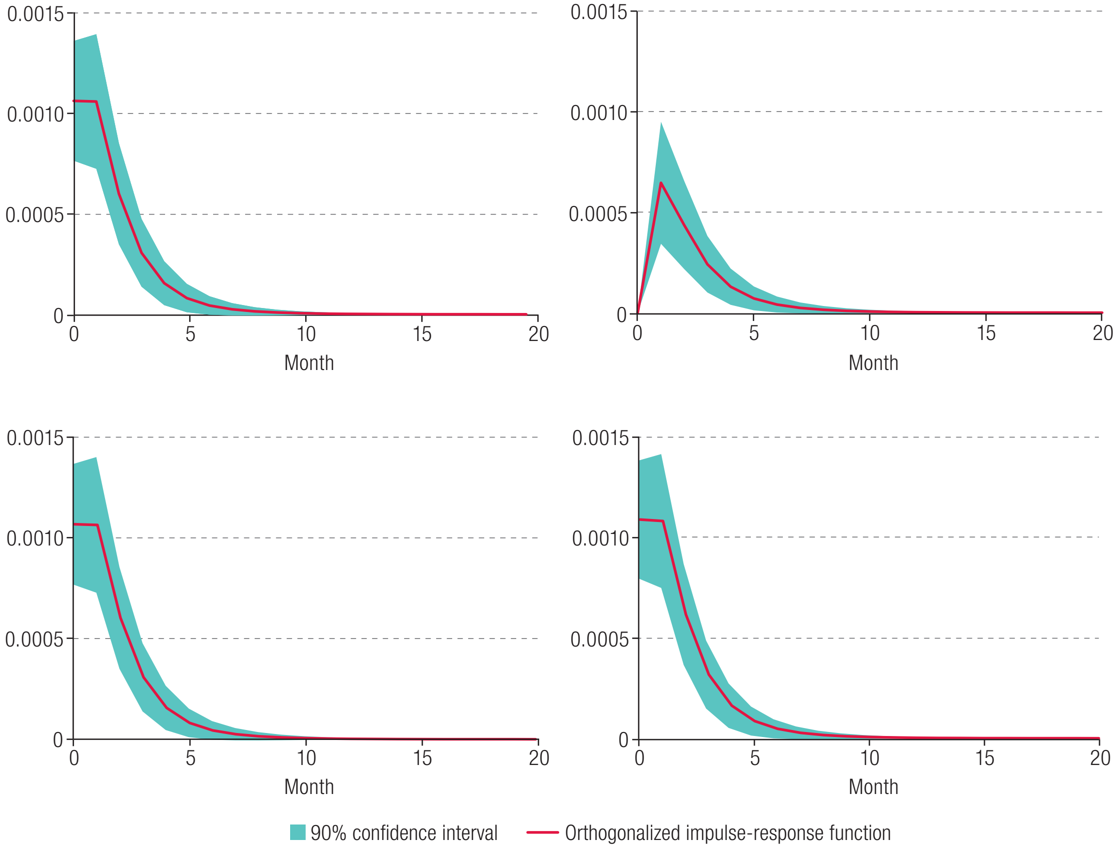


Figure 3 (concluded)



Source: Prepared by the authors.

In annex A1, the IRFs are presented without orthogonalizing. Figure A1.1 shows that the effect on inflation of a shock in the form of a 1% increase in the movement of the nominal exchange rate is positive and peaks at almost 0.03% in about the second month before falling back monotonically to zero by the seventh month. Figure A1.2 shows that an unexpected increase of 1% in the movement of the MPR is associated with an increase in inflation in the short term, peaking at approximately 0.003% in the second month, after which it begins to be statistically non-significant.

The results of the Granger causality test presented in table 7 show with 90% confidence that movements in the MPR and in the exchange rate (and in the two variables jointly) help predict inflation. Movements in the exchange rate (and in this jointly with inflation) help predict movements in the MPR. Since movement of the MPR is a control variable that is affected by expectations about relevant macroeconomic variables, including exchange-rate movements, this result is consistent and suggests that changes in MPR movements are associated with changes in future inflation.

The forecast error decomposition, meanwhile, indicates with 90% confidence that the variance of the inflation forecasting error is not affected by a shock in the MPR but is affected by movements in the exchange rate. In particular, the portion of the variance of the inflation forecasting error that is due to an exchange rate shock increases in the short run from 0% to approximately 22% after the fourth month.

Table 7
Granger causality test

Null hypothesis			Chi-squared	Probability > chi-squared
Inflation	does not cause	change in nominal exchange rate (NER)	0.71468	0.398
Change in monetary policy rate (MPR)	does not cause	change in NER	0.57277	0.449
Change in inflation and MPR jointly	does not cause	change in NER	1.7919	0.408
Change in NER	does not cause	inflation	12.533	0
Change in MPR	does not cause	inflation	2.8435	0.092
Change in NER and MPR jointly	does not cause	inflation	15.082	0.001
Change in NER	does not cause	change in MPR	4.6762	0.031
Inflation	does not cause	change in MPR	0.1271	0.721
Change in NER and inflation jointly	does not cause	change in MPR	4.7643	0.092

Source: Prepared by the authors.

VI. Conclusion

Three important conclusions emerge from this study. The main one is the corroboration that domestic demand has played a rather limited role as an inflationary factor in Chile. In normal periods, increases in domestic demand generally explain no more than 25% of observed inflation. The average monthly inflation observed during the 2000–2021 period was 0.3%, and of this we estimate that in normal periods demand growth explained inflation of 0.08% per month.

The second, surprising conclusion is that even the extraordinary periods of rapid acceleration in demand as a result of highly expansionary fiscal policies, AFP withdrawals or both had a rather modest effect on inflation. Only in the last 8 months of 2021 can we detect some effects from the expansion in demand on inflation, and even then it explained less than 28% of the increase in inflation that occurred in those months.

Thirdly and lastly, this study corroborates something that is to be expected in a small, open economy like Chile's: most domestic inflation is determined by foreign inflation.

The analysis in this paper is likely to be generally valid for small, open economies exposed to exogenous world prices, like those of most Latin American countries. It shows that, in these economies, demand stimulus may induce mainly output responses rather than price effects, as indeed happened in 2021, when gross domestic product increased by almost 11% in Chile.[6] In other words, these are the economies where Keynesian demand policies are likely to be most effective as an instrument for promoting greater employment and higher economic growth with little risk of inflation. Additionally, for small, open countries characterized by extreme levels of inequality, as most Latin American countries are, often in association with chronically inadequate demand as a binding constraint on economic growth, correcting this lack of demand may represent an opportunity to revive economic development through a virtuous spiral of faster economic growth, increasing use of installed capacity, higher investment and further output expansion.

Why do central banks in small, open economies so often insist on raising interest rates in their efforts to control domestic inflation fuelled by external factors? One possible explanation is that they expect tight monetary policy to strengthen the exchange rate. However, what we have shown in this paper is that domestic interest rates are not powerful enough to check exchange-rate depreciation

[6] This large output expansion was stifled in 2022 as a consequence of extremely restrictive monetary and fiscal policies implemented in the first three quarters of that year.

and that they therefore have only a small impact on domestic prices through this mechanism. Central banks come under pressure to “do something” when inflation accelerates, and the only instrument they have available (especially when other more direct measures such as exchange controls are out of the question for political reasons) are higher interest rates. The key point, however, is that raising rates in the face of external inflation does little to reduce domestic inflation, as this paper shows, but is very effective at damaging output. In other words, this policy is likely to cause stagflation.

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Annex A1

Descriptive statistics

Table A1.1 presents a summary of the data used in the estimates.

Table A1.1
Summary of the data used in the estimates

Variable	Number of observations	Mean	Standard error	Minimum	Maximum
Chile inflation	264	0.003	0.004	-0.012	0.015
United States inflation	264	0.002	0.004	-0.019	0.012
China inflation	264	0.002	0.006	-0.014	0.026
Eurozone inflation	264	0.001	0.004	-0.015	0.013
Change in nominal exchange rate	264	0.002	0.027	-0.070	0.166
Change in copper price	264	0.008	0.063	-0.295	0.260
Change in Federal Funds Effective Rate	264	0.002	0.159	-0.923	1.000
Dummy 1	264	0.049	0.217	0.000	1.000
Dummy 2	264	0.019	0.137	0.000	1.000
Dummy 3	264	0.015	0.122	0.000	1.000
Dummy 4	264	0.019	0.137	0.000	1.000
Dummy 5	264	0.015	0.122	0.000	1.000
Dummy 6	264	0.030	0.172	0.000	1.000

Source: Prepared by the authors.

Simulations

Table A1.2 presents a simulation of monthly inflation for 2021 and for January and February 2022, which are outside the sample. As can be seen, the simulations based on the estimated coefficients replicate the inflation rates reasonably well in most months.

Table A1.2
Simulated contributions to monthly inflation estimated by two-stage least squares, January 2021 to February 2022[a]
(Percentages)

Year	Month	Monthly inflation	Percentage change in nominal exchange rate	United States inflation	China inflation	Eurozone inflation	Chile inflation with a one-period lag	Simulation: predicted value
2021	1	0.7	-0.091	0.191	0.027	0.007	0.111	0.245
2021	2	0.2	-0.008	0.245	0.019	0.007	0.259	0.523
2021	3	0.4	0.031	0.317	-0.014	0.033	0.074	0.442
2021	4	0.4	-0.153	0.368	-0.011	0.022	0.148	0.374
2021	5	0.3	0.037	0.359	-0.003	0.011	0.148	0.553
2021	6	0.1	0.121	0.416	-0.011	0.011	0.111	0.648
2021	7	0.8	0.198	0.216	0.008	-0.004	0.037	0.455
2021	8	0.4	0.235	0.093	0.003	0.015	0.296	0.641
2021	9	1.2	0.029	0.122	0.000	0.019	0.148	0.317
2021	10	1.3	0.233	0.372	0.022	0.030	0.443	1.100
2021	11	0.5	-0.010	0.220	0.011	0.015	0.480	0.716

Table A1.2 (concluded)

Year	Month	Monthly inflation	Percentage change in nominal exchange rate	United States inflation	China inflation	Eurozone inflation	Chile inflation with a one-period lag	Simulation: predicted value
2021	12	0.8	0.270	0.138	-0.008	0.015	0.185	0.599
2022	1	1.2	-0.192	0.358	0.012	0.011	0.296	0.486
2022	2	0.3	-0.108	0.403	0.018	0.033	0.443	0.790

Source: Prepared by the authors.
[a] January and February 2022 are outside the sample.

Robustness tests

Several model robustness tests were implemented. First, the Monthly Index of Economic Activity (IMACEC) was considered as an additional control variable, serving as a proxy for demand effects. Table A1.3 shows these estimators. As can be seen, the fundamental results (domestic inflation is explained primarily by foreign inflation and the exchange rate) are maintained. The role of demand factors in both normal and exceptional periods is of minor importance.

Table A1.3
Second stage of the two-stage least squares estimation including changes in the Monthly Index of Economic Activity (IMACEC)

Number of observations	263
Wald chi-squared (12)	205.60
Probability > chi-squared	0.0000
R-squared	0.4373

Chile inflation	Coefficient	Robust standard error	z	$p > \|z\|$	Confidence interval (95%)	
Percentage change in nominal exchange rate	0.0622	0.0134	4.65	0.000	0.0360	0.0884
United States inflation	0.4722	0.0545	8.66	0.000	0.3653	0.5790
China inflation	0.0443	0.0304	1.46	0.144	-0.0152	0.1038
Eurozone inflation	-0.0072	0.0550	-0.13	0.896	-0.1149	0.1005
Dummy 1	-0.0005	0.0014	-0.38	0.704	-0.0032	0.0021
Dummy 2	-0.0014	0.0014	-1.01	0.315	-0.0042	0.0014
Dummy 3	0.0000	0.0008	0.03	0.974	-0.0016	0.0016
Dummy 4	0.0014	0.0014	0.99	0.320	-0.0014	0.0042
Dummy 5	-0.0002	0.0016	-0.14	0.888	-0.0033	0.0029
Dummy 6	-0.0004	0.0016	-0.22	0.827	-0.0035	0.0028
Chile inflation with a one-period lag	0.3707	0.0565	6.56	0.000	0.2600	0.4813
Percentage change in IMACEC	0.0057	0.0041	1.40	0.161	-0.0023	0.0136
γ_0	0.0007	0.0003	2.50	0.012	0.0001	0.0012

Source: Prepared by the authors.

When we use IMACEC as a proxy for demand stimuli and compare the results of table A1.3 with those of table 1 in the body of the text, the results are practically identical. The estimated coefficients remain very stable; specifically, IMACEC does not seem to be a variable of importance in explaining inflation. The same holds when the GMM method is used in the estimation (see table A1.4).

Table A1.4
Estimation by the generalized method of moments including changes in the Monthly Index of Economic Activity (IMACEC)

Number of observations	263
Wald chi-squared (12)	208.80
Probability > chi-squared	0.0000
R-squared	0.4373

Chile inflation	Coefficient	Robust standard error	z	$p > \|z$	Confidence interval (95%)	
Percentage change in nominal exchange rate	0.0622	0.0134	4.66	0.000	0.0361	0.0884
United States inflation	0.4724	0.0545	8.67	0.000	0.3656	0.5791
China inflation	0.0442	0.0303	1.46	0.145	-0.0153	0.1036
Eurozone inflation	-0.0067	0.0547	-0.12	0.903	-0.1138	0.1005
Dummy 1	-0.0005	0.0013	-0.39	0.695	-0.0032	0.0021
Dummy 2	-0.0014	0.0014	-1.02	0.310	-0.0042	0.0013
Dummy 3	0.0000	0.0008	0.03	0.975	-0.0016	0.0016
Dummy 4	0.0014	0.0014	1.00	0.319	-0.0014	0.0042
Dummy 5	-0.0002	0.0016	-0.14	0.888	-0.0033	0.0029
Dummy 6	-0.0004	0.0016	-0.22	0.827	-0.0035	0.0028
Chile inflation with a one-period lag	0.3708	0.0564	6.57	0.000	0.2602	0.4814
Percentage change in IMACEC	0.0056	0.0040	1.40	0.161	-0.0022	0.0135
γ_0	0.0006	0.0003	2.50	0.013	0.0001	0.0012

Source: Prepared by the authors.

Another test of the robustness of the estimators is implemented using the monetary policy rate (MPR) directly as the explanatory variable for monthly inflation. Since the effect of the MPR on inflation is expected to be subject to lags, we use this variable with three lags. Table A1.5 shows these results. As can be seen there, the inclusion of the lagged MPR variables does not affect the value of the key coefficients, which continue to be external inflation and exchange-rate movements. Surprisingly, the MPR does not appear to be statistically significant in any of its lags. The different lags of the MPR are also non-significant when we estimate the first stage of the estimation by TSLS (see table A1.6).

Table A1.5
Second stage of the estimation by two-stage least squares including changes in the monetary policy rate (MPR)

Number of observations	261
Wald chi-squared (14)	228.42
Probability > chi-squared	0.0000
R-squared	0.4434

Chile inflation	Coefficient	Robust standard error	z	$p > \|z\|$	Confidence interval (95%)	
Percentage change in nominal exchange rate	0.0576	0.0129	4.470	0.000	0.0324	0.0828
United States inflation	0.4488	0.0559	8.020	0.000	0.3392	0.5584
China inflation	0.0201	0.0312	0.640	0.520	-0.0411	0.0812
Eurozone inflation	0.0302	0.0428	0.710	0.480	-0.0537	0.1141

Table A1.5 (concluded)

Chile inflation	Coefficient	Robust standard error	z	$p > \|z\|$	Confidence interval (95%)	
Dummy 1	-0.0002	0.0014	-0.130	0.896	-0.0029	0.0025
Dummy 2	-0.0011	0.0013	-0.850	0.395	-0.0038	0.0015
Dummy 3	0.0002	0.0010	0.200	0.844	-0.0018	0.0022
Dummy 4	0.0017	0.0013	1.270	0.203	-0.0009	0.0043
Dummy 5	-0.0004	0.0014	-0.250	0.802	-0.0032	0.0025
Dummy 6	-0.0008	0.0018	-0.430	0.667	-0.0042	0.0027
Chile inflation with a one-period lag	0.3639	0.0593	6.130	0.000	0.2476	0.4801
Percentage change in MPR (one-period lag)	0.0025	0.0019	1.310	0.189	-0.0012	0.0061
Percentage change in MPR (two-period lag)	-0.0018	0.0023	-0.780	0.434	-0.0063	0.0027
Percentage change in MPR (three-period lag)	0.0023	0.0014	1.620	0.105	-0.0005	0.0051
γ_0	0.0007	0.0003	2.680	0.007	0.0002	0.0012

Source: Prepared by the authors.

Table A1.6
First stage of the estimation by two-stage least squares including changes in the monetary policy rate (MPR)

Number of observations	261
F (15, 245)	6.90
Probability > F	0.0000
R-squared	0.3405
Adjusted R-squared	0.3001

Percentage change in nominal exchange rate	Coefficient	Robust standard error	t	$p > \|t\|$	Confidence interval (95%)	
United States inflation	0.0953	0.4582	0.21	0.835	-0.8072	0.9978
China inflation	-0.3025	0.2429	-1.25	0.214	-0.7810	0.1760
Eurozone inflation	0.1039	0.3347	0.31	0.757	-0.5553	0.7631
Dummy 1	-0.0046	0.0110	-0.42	0.678	-0.0263	0.0171
Dummy 2	-0.0039	0.0169	-0.23	0.818	-0.0372	0.0295
Dummy 3	-0.0162	0.0110	-1.48	0.141	-0.0378	0.0054
Dummy 4	-0.0079	0.0075	-1.05	0.293	-0.0226	0.0068
Dummy 5	-0.0055	0.0078	-0.71	0.480	-0.0208	0.0098
Dummy 6	0.0240	0.0067	3.59	0.000	0.0108	0.0372
Chile inflation with a one-period lag	0.1934	0.4439	0.44	0.663	-0.6808	1.0677
Percentage change in MPR (one-period lag)	0.0012	0.0135	0.09	0.929	-0.0253	0.0277
Percentage change in MPR (two-period lag)	-0.0265	0.0178	-1.48	0.139	-0.0616	0.0087
Percentage change in MPR (three-period lag)	0.0039	0.0153	0.25	0.801	-0.0263	0.0340
Percentage change in copper price	-0.2139	0.0382	-5.61	0.000	-0.2890	-0.1387
Percentage change in Federal Funds Effective Rate	-0.0185	0.0083	-2.21	0.028	-0.0349	-0.0020
γ_0	0.0041	0.0019	2.09	0.038	0.0002	0.0079

Source: Prepared by the authors.

Additionally, the MPR is not statistically significant when the GMM method is used (see table A1.7).

Table A1.7
Estimation by the generalized method of moments including changes in the monetary policy rate (MPR)

Number of observations	261
Wald chi-squared (14)	229.85
Probability > chi-squared	0.0000
R-squared	0.4434

Chile inflation	Coefficient	Robust standard error	z	$p > \|z\|$	Confidence interval (95%)	
Percentage change in nominal exchange rate	0.0576	0.0129	4.47	0.000	0.0323	0.0828
United States inflation	0.4486	0.0557	8.05	0.000	0.3394	0.5579
China inflation	0.0201	0.0312	0.64	0.519	-0.0410	0.0812
Eurozone inflation	0.0302	0.0428	0.70	0.481	-0.0537	0.1140
Dummy 1	-0.0002	0.0014	-0.13	0.900	-0.0028	0.0025
Dummy 2	-0.0011	0.0013	-0.85	0.395	-0.0037	0.0015
Dummy 3	0.0002	0.0010	0.19	0.847	-0.0018	0.0022
Dummy 4	0.0017	0.0013	1.27	0.203	-0.0009	0.0043
Dummy 5	-0.0004	0.0014	-0.25	0.802	-0.0032	0.0025
Dummy 6	-0.0008	0.0018	-0.43	0.668	-0.0042	0.0027
Chile inflation with a one-period lag	0.3637	0.0592	6.15	0.000	0.2477	0.4797
Percentage change in MPR (one-period lag)	0.0025	0.0019	1.32	0.188	-0.0012	0.0061
Percentage change in MPR (two-period lag)	-0.0018	0.0023	-0.78	0.434	-0.0063	0.0027
Percentage change in MPR (three-period lag)	0.0023	0.0014	1.62	0.105	-0.0005	0.0051
γ_0	0.0007	0.0003	2.69	0.007	0.0002	0.0012

Source: Prepared by the authors.

Vector autoregression model: impulse-response functions

Figure A1.1
Impulse-response function for the effect of an exchange-rate shock of a 1% increase on inflation for different orders in the Cholesky decomposition

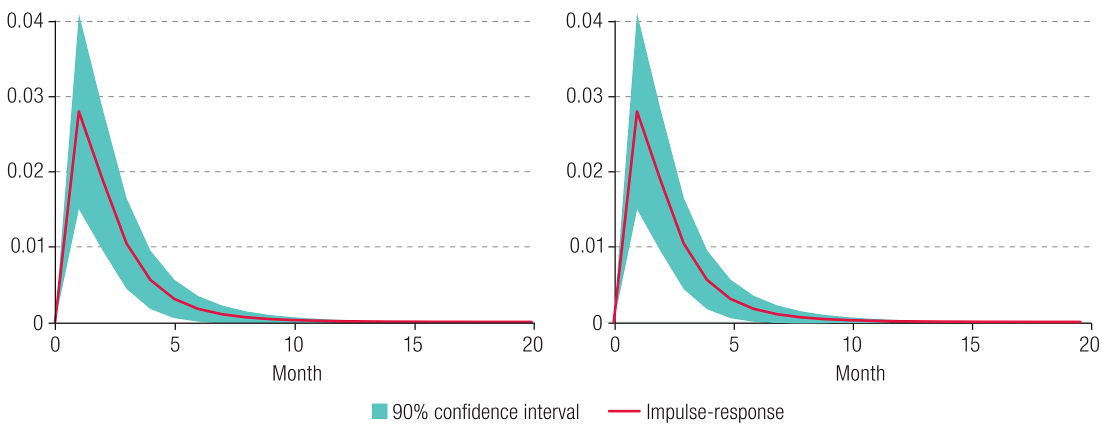
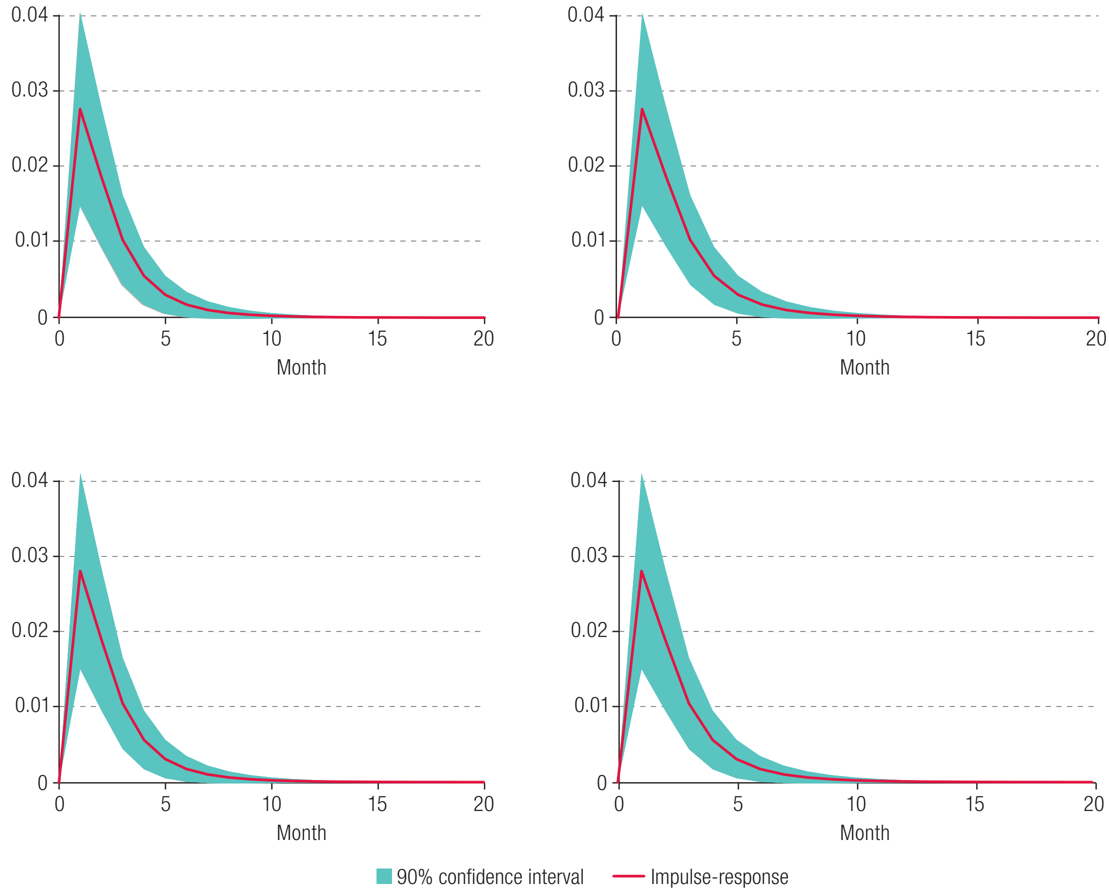


Figure A1.1 (concluded)



Source: Prepared by the authors.

Figure A1.2
Impulse-response function for the effect of a monetary policy rate shock of a 1% increase on inflation for different orders in the Cholesky decomposition

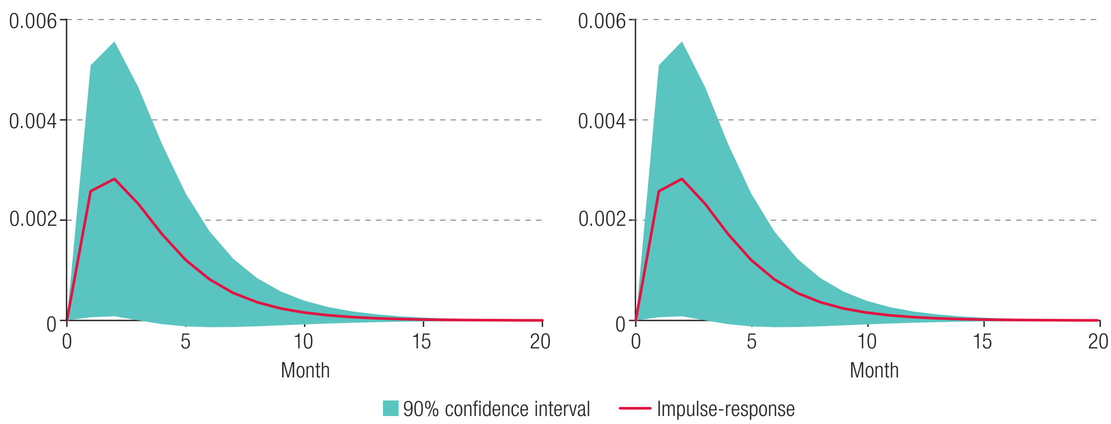
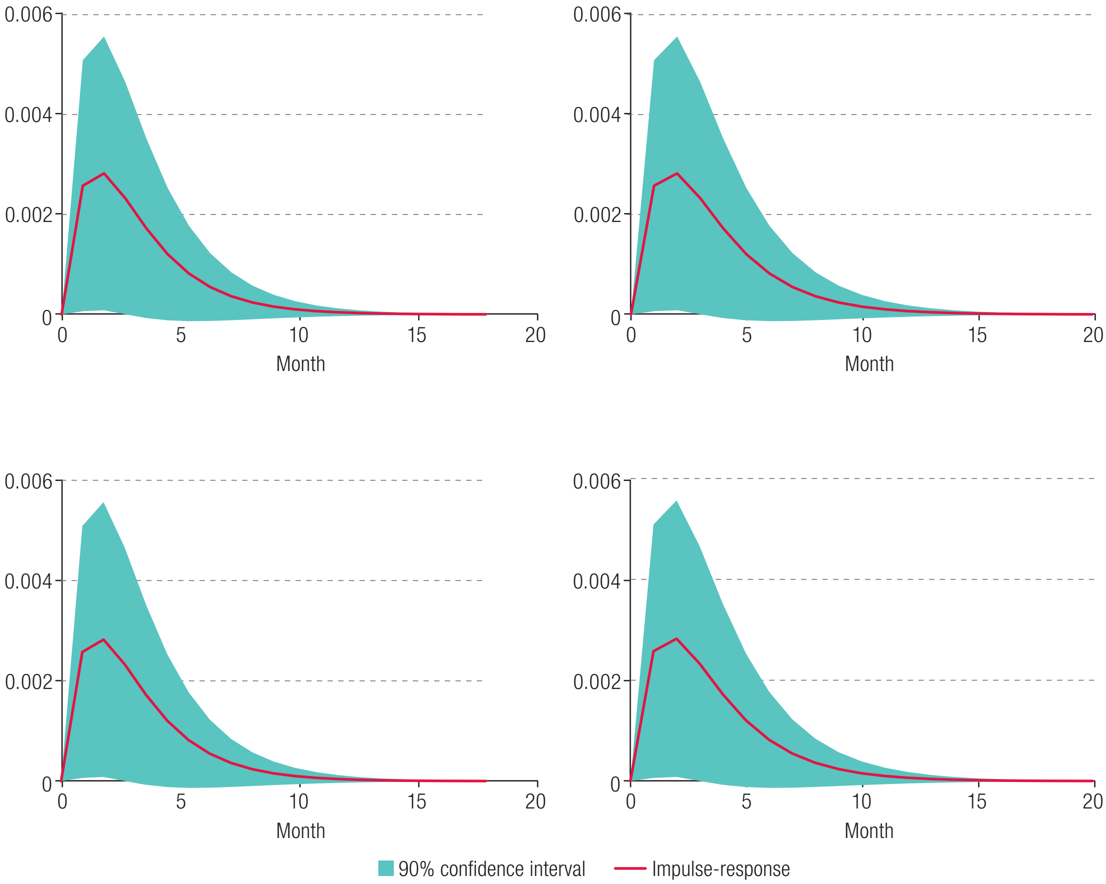


Figure A1.2 (concluded)



Source: Prepared by the authors.