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informes y estudios especiales

Returning to an eternal debate: the terms of trade for commodities in the twentieth century

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Contents

Abstract

This paper looks at the evolution of the terms of trade between commodities and manufactures in the twentieth century. A statistical analysis of the relative price series for 24 commodities and of eight indices reveals a significant deterioration in their barter terms of trade over the course of the twentieth century. This decline was neither continuous, nor was it distributed evenly among individual products, however. The data show that the far-reaching changes that the world economy underwent around 1920 and again around 1980 led to a stepwise deterioration which, over the long term, was reflected in a decline of nearly 1% per year in aggregate real prices for raw materials.

JEL classification: F12, 013, Q00

Introduction

This paper examines the evolution of the international terms of trade for commodities in the light of recent empirical evidence. The analysis is divided into nine sections. The second provides a brief summary of the theoretical foundations for what is known as the Prebisch-Singer hypothesis. The third presents the evidence, in the most straightforward manner possible, of the evolution of relative prices (or barter terms of trade) for commodities as compared to manufactures. The fourth section provides an analysis of the autoregressive dynamics of the series, while the fifth assesses the possible existence of structural breaks. In the sixth section, the dynamics of the series are examined by drawing upon the analyses set forth in the preceding two sections. The seventh looks at the evolution of the stationary series and identifies two points in time when major changes occurred (1921 and 1979). The eighth provides a brief discussion of the persistence of shocks in the short and medium terms. The study's findings are interpreted in the final section.

The empirical analysis employs price series for 24 commodities, seven indices constructed by Grilli and Yang (1988) and, as an alternative index, the industrial commodities price index used by *The Economist*. In order to cover the entire twentieth century, the indices developed by Grilli and Yang have been updated to 2000 using these authors' original methodology. A detailed description of the products and indices is provided in section III.

I. The Prebisch-Singer hypothesis

The thesis concerning the declining trends of the terms of trade for developing countries was formulated concurrently by Sir Hans Singer (1950) and Raul Prebisch (1951) in the early 1950s. Their work in this area was, in large part, undertaken in an effort to account for the empirical research findings of the United Nations Department of Economic and Social Affairs, which corroborated the existence of such a trend.[1] The original formulation of this thesis combined two different yet clearly complementary hypotheses whose subsequent theoretical development in the economic literature was to follow parallel courses. One of these hypotheses regarded the negative effect of the income-inelasticity of the demand for commodities on the developing countries' terms of trade and, the other, the asymmetries in the functioning of labour markets in the world economy's "centre" and "periphery". The fundamental conceptual difference between these two hypotheses lies in the fact that, whereas in the first case, the pressure towards a deterioration in real commodity prices is generated in goods markets (i.e., on the *barter* terms of trade), in the second this pressure is generated in factor markets (and hence on the *factorial* terms of trade) and thus affects the barter terms of trade only indirectly, through the effects on production costs. Another difference, which is a consequence of the above, is that, whereas the first hypothesis applies solely to commodities (or, more generally, to goods whose demand exhibits a low income-elasticity), the second applies to all goods and services produced in developing countries, regardless of what types of goods or services they are or the nature of final demand.

1 The relevance and reliability of the data that were used have been debated at length. See, in particular, Scandizzo and Diakosawas (1987).

The first hypothesis was based on the well-known observation that economic growth tends to trigger changes in the production structure over time and, in particular, to generate a tendency towards a relative reduction in the size of the primary sector. As is widely recognized, this structural break is associated not only with the characteristics of final demand (especially the low income-elasticity of the demand for foodstuffs) but also with the fact that, in many cases, technological change in the production of manufactures entails reductions in raw materials costs or the production of synthetic materials. These variations in the production structure have important implications at the world level if the international division of labour is such that developing countries specialize in the production of raw materials while industrialized nations specialize in manufactures. Under these circumstances, it is to be expected that either the former will grow more slowly or the surplus primary commodities which they produce will tend to push down the relative international prices of those commodities.[2]

The second hypothesis was formulated by both authors —although perhaps more clearly by Singer— in terms of an unequal distribution of the fruits of technological progress. According to this hypothesis, in the case of manufactures, these benefits are distributed to producers in the form of higher income, but in the case of commodities they are reflected in lower prices. This asymmetry is a result of the way in which both goods markets (greater market power for setting the prices of manufactures) and labour markets (greater organization of industrial workers) operate. At the international level, however, it is also a reflection of the international division of labour. In this instance, the more precise formulation was made by Prebisch. He thought that, because of the weaker long-term demand for raw materials, the relative surplus of labour displaced from primary activities tends to concentrate in developing countries, which, in turn, have more difficulty putting that surplus labour to work in new production sectors. The problems they face include political restrictions on migration to industrialized nations and the obstacles hindering late industrialization, which, in their view, are associated with the striking disparities between the countries of the "centre" and the "periphery" in terms of technological capabilities and the availability of capital. This situation generates a surplus of labour which leads to a relative decline in the wages of developing-country workers and, hence, in those countries' terms of trade.[3]

The history of the controversy surrounding the issue of developing countries' terms of trade can largely be written by tracing the development of these two hypotheses (see Ocampo, 1986 and 1993). The neoclassical and Keynesian literature of the 1950s and 1960s focused on the first of these mechanisms. According to Johnson (1954), the lower income-elasticity of the demand for raw materials ought to be reflected in slower economic growth in the countries specializing in those products or in a tendency for raw material prices to decline. This effect depends entirely on *income*-elasticity, but the lower the *price*-elasticity of the demand for raw materials is, the larger the decrease will be. It should be emphasized, however, that this type of model is incapable of generating asymmetries in the transmission of technical progress, and thus cannot be used to validate the second Prebisch-Singer hypothesis.

2 The pressure towards unequal rates of growth will be greater if the externalities generated by production (the generation of demand multipliers and the externalities associated with technical progress, in particular) are greater in the case of industrial production. This was also one of the core elements of the two authors' thesis, but will not be discussed further here. For an interesting recent empirical assessment of this question covering the period 1870-1940, see Hadaas and Williamson (2001).

3 Prebisch believed that this asymmetry was particularly evident during downswings in the business cycle. In his view, workers in countries of the centre were not only able to secure wage increases during booms, but were also able to defend their wages during recessionary phases in the world business cycle. In contrast, because of the surplus —and, hence, the marked cyclical deterioration in raw material prices— workers in the periphery were unable to prevent the deterioration of their income levels during crises.

In a neoclassical (Heckscher-Ohlin) trade model, any factor that increases the supply of a given good will result in a decrease in its relative price. Thus, in countries or regions that are large enough to influence international prices, technological change in export industries will be reflected in a deterioration in the barter terms of trade. On the other hand, technological change in import-substitution industries will have just the opposite effect, since it will cause factors of production to be transferred to those sectors, thereby reducing the supply of exports and, consequently, improving the terms of trade.

Unlike this line of reasoning, the analysis of "unequal exchange" since the late 1960s focused on asymmetries in the operation of labour markets. The most comprehensive treatments of this subject are based on the models developed by Findlay (1980 and 1981) and Taylor (1983, chapter 10) in the early 1980s (for a comparison of these and other models, see Ocampo, 1986). In both cases, the economy being modelled is one in which the "North" determines the pace of the world economy's growth and in which the "South" adapts to that pace. The essential element of this model is, however, its recognition of the asymmetries in the economic structures of both. Thus, the North has a neoclassical economic structure in Findlay's model and a Keynesian (or, more accurately, Kaleckian) structure in Taylor's, while in both formulations the South functions as a Lewis-type surplus-labour economy. These asymmetries give rise to a pattern that fits in perfectly with the second Prebisch-Singer hypothesis, i.e., in the long run, the North appropriates the full benefits of its own process of technical change, while the South's productivity gains lead to a commensurate deterioration in its barter terms of trade (in other words, its technical changes are "exported"). This is a reflection of the asymmetrical effects that technological change has on real wages. While in the North, wage increases are proportional to increases in productivity, in the South real wages are not affected by technological change. The corresponding effect is transmitted through production costs and is therefore unrelated to the type of good being produced or the demand for it.[4]

Over the last two decades, the empirical literature on this issue has been greatly enriched,[5] thanks to the existence of more reliable data and new, more rigorous statistical methodologies for analysing time series. In the following section, these methodologies are used to analyse the validity of a thesis that has already been around for half a century. It should be added that, given available data, the empirical assessment refers chiefly to the barter terms of trade. Nevertheless, the abundant recent literature on the international "convergence" or "divergence" of per capita incomes and wages can, in a sense, be regarded as a contribution to the clarification of the second above-mentioned hypothesis regarding the trend in the factorial terms of trade.[6]

4 Consequently, contrary to the argument made by Hadaas and Williamson (2001), this effect should be modelled under the assumption of equivalent (unitary) income elasticities for goods produced by both regions.

5 See, among others, Cuddington and Urzúa (1989), Powell (1991), Ardeni and Wright (1992), Cuddington (1992), Cuddington and Wei (1998), Bleany and Greenaway (1993), León and Soto (1995a and 1995b), Cashin and McDermott (2002) and Cuddington, Ludema and Jayasuriya (2002).

6 For a summary of the conclusions reached in the course of this debate and some additional calculations, see ECLAC (2002, chapter 3).

II. Preliminary analysis of the trend in real commodity prices

This analysis is based on 24 commodity price series[7] and seven indices which were originally developed by Grilli and Yang (1988) for the period 1900-1986 but which have been updated to the year 2000 in order to cover the entire century.[8] The deflator used to calculate real prices is the Manufacturing Unit Value (MUV) index developed by the United Nations.[9] As an alternative, the industrial commodity price index of The Economist for 1880-1999 is also used.[10] The deflator used in this case was the Great Britain Index of Export Prices[11] for 1880-1900 and the MUV thereafter. Figure 1A shows the total Grilli and Yang price indices. The nomenclature used in the rest of the paper is explained in box 1.

7 The products used include six metals (aluminium, copper, tin, silver, lead and zinc), six non-food raw materials (palm oil, cotton, rubber, leather, wool, timber and jute), seven food products (rice, sugar, bananas, lamb, beef, maize and wheat), three beverages (cocoa, coffee and tea) and tobacco. John Cuddington of Georgetown University was kind enough to make these series available to the authors.

8 The indices covering the period from 1986 to 2000 and the methodology used to update them are presented in appendix I.

9 This index reflects the unit value of industrialized countries' exports of manufactured products. It was originally taken from Grilli and Yang (1988) and later updated with data compiled by the United Nations.

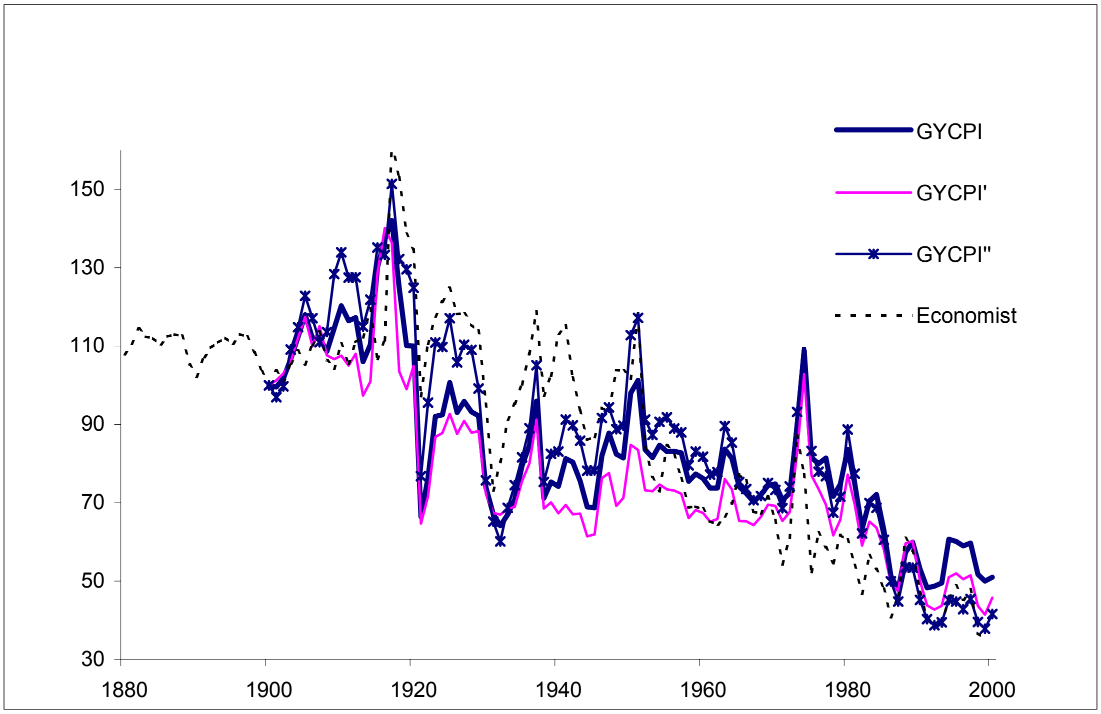
10 This series includes prices for six metals (aluminium, copper, nickel, zinc, tin and lead) and nine non-agricultural commodities (cotton, timber, leather, rubber, wool, palm oil, soybeans and soybean oil). The authors were able to obtain this series thanks to the generous assistance of Mr. Paul Cashin of the Research Department of the International Monetary Fund.

11 This index (developed by A.G. Silverman) was chosen due to the absence of alternative data and the importance of this country in world trade in the late nineteenth century.

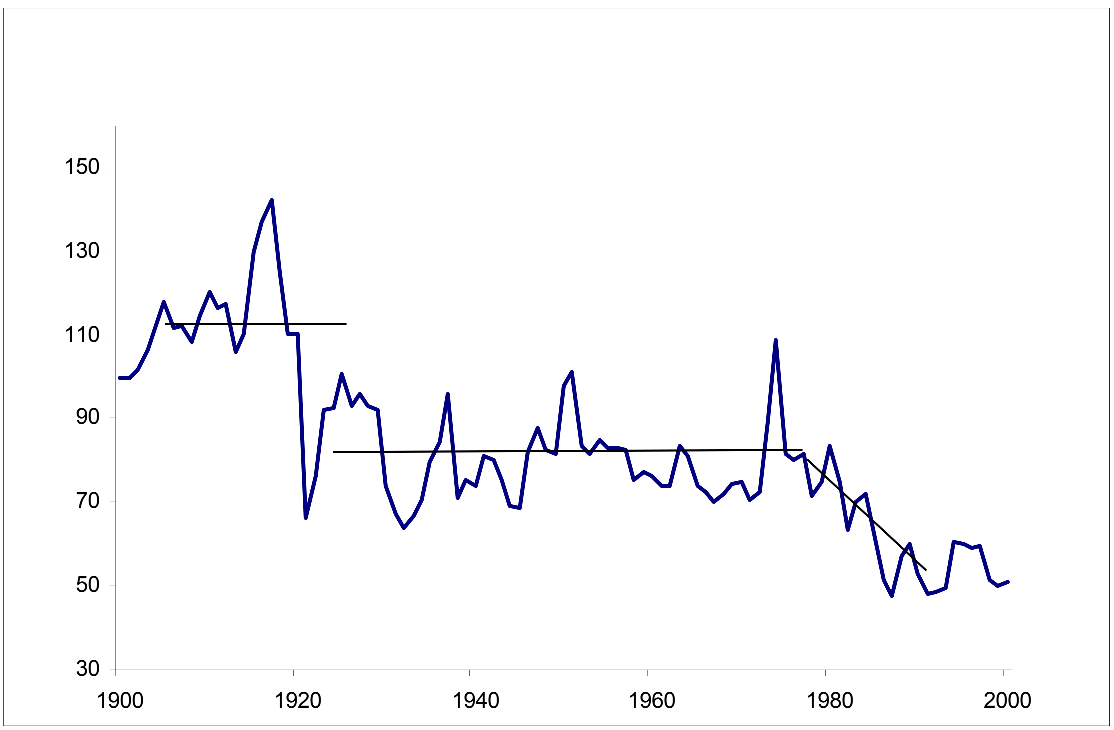
Figure 1

TOTAL NON-FUEL COMMODITY PRICE INDICES, 1900=100

A. The Economist and the Grilli and Yang indices



B. Total non-fuel commodity price index (GYCPI)



Source: Grilli and Yang (1988); *The Economist* and calculations based on United Nations data.

Box 1

NOMENCLATURE

Using the nomenclature employed by Grilli and Yang, seven dollar-denominated price indices for tradable commodities will be considered:

GYCPI: Total index, weighted by the share of total exports represented by each product in 1977-1979; three subindices are also derived: *food products, non-food products* and *metals*.

GYCPI': Total index, weighted by the developing countries' share of commodity exports in 1981. (The original index used weightings for 1977-1979; since these weightings were unavailable, weightings for 1981 were substituted.)

GYCPI'': Total index, weighted by the share of world exports represented by commodities during the year in question.

GYCPI''': The same as GYCPI'' except that it also includes oil prices.

Source: Grilli and Yang (1988).

Despite the differences between the series caused by the varying weightings of the products' shares of total export and the series' cyclical variability, a marked long-term downturn is clearly seen. Taken together, the cumulative decrease is very large, since between 1900 and 2000, raw materials lost between 50% and 60% of their value relative to manufactures. This finding has been corroborated by various authors.[12] Cashin and McDermott (2002), for example, found a downward trend without structural breaks of 1.3% per year over a period of 140 years,[13] which, oddly enough, they interpret as small compared to the variability of prices, even though it translates into a cumulative decline of 75% over the period they analysed.

This downturn is a hallmark of the *twentieth century*, not the nineteenth. In fact, in keeping with the recent observations of Hadaas and Williamson (2001), the series actually point to an improvement in real raw material prices in the late nineteenth and early twentieth centuries.[14] These authors have also said that the sharp reduction in shipping costs that occurred during those years benefited all countries. This is reflected in the improvement that can be seen in the terms of trade for the period when prices are measured in a given location (i.e., f.o.b. export prices versus c.i.f. import prices).

This decrease has not been continuous. Instead, it has occurred in stepwise shifts that appear to have permanently altered price levels. Figure 1B graphs the GYCPI series, highlighting what appear to be these shifts. It is noteworthy that the largest price drops followed, with a lag, the two major slowdowns in the industrialized economies' long-term growth rates during the First World War and in 1973, respectively (see Madisson, 1995).

These observations suggest that, rather than discussing whether or not there was a long-term downtrend in the barter terms of trade for raw materials during the twentieth century, it would be more to the point to talk about the particular dynamics exhibited by this decline and how the evolution of prices of individual products differ. The Prebisch-Singer hypothesis has traditionally *—and perhaps erroneously—*[15] been associated with a secular or continuous trend. This study considers the hypothesis that this deterioration occurred in steps. The exploration of this possibility is based on a detailed analysis of the behaviour of aggregate price indices and price indices for individual products.

A preliminary view of the situation is outlined in table 1, which shows the series' average annual growth rates for the two stages during which the sharpest decreases were seen, for the period as

12 See, for example, a summary of the studies conducted up to the 1980s in Ocampo (1993).

13 Cashing and McDermott use *The Economist*'s industrial commodity price series for the period 1862-1999 and then convert it to relative prices using the United States GDP as a deflator.

14 A number of country studies also provide evidence that raw material prices rose in real terms throughout the nineteenth century.

15 Cuddington et al. (2002) contend that the Prebisch-Singer hypothesis did not say that the long-term trend was necessarily constant over time, but only that it was negative.

a whole and calculations of the loss in relative value that occurred between the first and last five years of the twentieth century. The data show that product behaviour was quite heterogeneous. However, regardless of the weighting used, all the indices show average declines of 0.8% per year due to the decrease in the relative value of food products, which was particularly steep during the 1920s and the 1980s. While non-food products lost around 15% of their value in the course of the century, food products lost half their purchasing power. The only commodities whose relative value rose substantially during the century as a whole were beef, lamb, timber and tobacco.

Table 1

COMMODITY PRICES AND INDICES DEFLATED BY THE MUV

(Average annual growth rates, percentages)

	1920-1930	1980-1990	1900-2000	1900/04-1996/200	
				Annual	Accumulated
Aluminium	1.8	2.8	-1.1	-1.3	-71.7
Bananas	5.8	0.1	0.0	-0.1	-7.5
Beef	-0.2	-6.6	1.0	0.9	134.6
Cocoa	-0.7	-9.5	-1.3	-1.0	-61.8
Coffee	0.4	-8.3	-0.1	0.4	45.3
Copper	1.4	-1.1	-0.7	-0.6	-46.0
Cotton	-3.0	-6.1	-1.0	-1.1	-66.0
Jute	-0.9	0.6	-0.4	-0.7	-30.4
Lamb	-0.1	-3.9	1.6	1.7	399.3
Lead	0.7	-4.2	-0.8	-0.3	-48.0
Leather	-4.7	1.3	-0.8	-1.1	-63.6
Maize	-1.2	-5.3	-0.8	1.2	-61.9
Palm oil	-2.8	-2.8	-0.3	0.0	-1.3
Rice	3.7	-6.9	-1.3	-1.2	-66.9
Rubber	-9.5	-7.9	-2.8	-2.8	-93.4
Silver	-5.3	-16.2	-0.3	-1.0	-23.8
Sugar	-16.8	-10.5	-1.3	-1.1	-65.4
Tea	7.6	-4.0	-0.7	0.7	-56.2
Timber	-2.2	-1.5	1.1	-1.5	208.1
Tin	0.1	-10.2	0.1	0.2	15.4
Tobacco	-2.7	-0.4	0.8	-0.7	100.4
Wheat	-4.5	-3.1	-0.6	-0.9	-46.4
Wool	-3.1	-5.5	-1.2	-0.4	-76.6
Zinc	-0.9	4.7	0.3	0.1	5.9
Indices					
GYCPI	-3.9	-4.4	-0.7	-0.7	-47.8
GYCPI’	-3.7	-4.2	-0.8	-0.8	-55.4
GYCPI”	-4.9	-6.5	-0.9	-1.0	-60.2
GYCPI’”	-5.1	-6.1	-0.4	-0.7	-49.3
Food products	-5.2	-7.8	-0.8	-0.7	-49.8
Non-food products	1.2	5.1	0.0	-0.2	-14.6
Metals	5.5	0.9	-0.1	-0.1	-7.1
Economist	-3.4	-2.5	-1.0	-1.0	-60.1

Source: Calculations based on data from Grilli and Yang (1988), *The Economist* and United Nations.

The following analysis deals with the dynamic structure of each series. While it is clear that during the twentieth century most commodities’ purchasing power dropped significantly, neither the size of this cumulative decrease nor the average annual growth rates provide a basis from which to infer the series’ long-term growth behaviour.[16] In order to understand their behaviour, it is necessary to know how the series will respond to shocks and, in particular, to know whether or not the series have a stochastic component. It is also important to evaluate the possibility that the series exhibit structural breaks before using all the information that has been compiled to model the behaviour of each one as completely as possible.

[16] As Cuddington et al. has noted, modern time series econometrics has taught us that it is potentially misleading to assess the presence of long-term trends by eyeballing the series or estimating simple log-linear time trend models.

III. Autoregressive dynamics of the series

A differentiation should be drawn between two types of autoregressive processes that may give rise to statistical trends that display different dynamics: a deterministic trend (DT), if the series are stationary in variance; and a stochastic trend (ST) in the case of series exhibiting non-stationarity in variance.[17]

A deterministic trend (DT) model exhibits the following dynamics:

$$\textbf{(v)} \qquad Log\ P_t = \beta\ T_t + ARMA(p,q)e_t$$

where Tt is a trend variable, et is an i.i.d. (independent and identically distributed) random shock and the parameter β is the trend (exponential growth rate), which can be estimated using traditional econometric procedures (ordinary least squares). The use of the term ARMA for the residual rules out the possibility of a misspecification caused by higher order autocorrelations of the series. In this model, the series Pt is not stationary (unless β=0), but the fluctuations of Pt around its DT are stationary (there is no evidence of a unit root). The only information required in this model in order to forecast the long-term price trend is the average growth rate of the variable (β), since, because they are wholly transitory, shocks will not affect long-term projections.

In turn, a stochastic or stationary trend (ST) model in differences has the following dynamics:

[17] See, among others, León and Soto (1995a) and Cuddington, et al. (2002).

$$\textbf{(V)} \qquad \Delta\, Log\, P_t = \gamma + ARMA(p', q')\, \mu_t$$

where Δ is the first differences operator and γ is the average growth rate of the variable. The presence of μ_t,[18] an i.i.d. random variable, will induce stochastic behaviour in price levels. This model would be appropriate if the series is found to have a unit root. Consequently, in addition to a possible deterministic trend (γ), in this case shocks can have permanent effects on commodity price levels. If γ is statistically significant, then we have a unit root process with drift.

Applying this logic to the GYCPI index, Cuddington et al. (2002) show that the possibility of finding a statistically significant trend hinges upon the conclusions derived from a unit root test. More specifically, these authors show that, if the index is found to be following a DT process, then the trend can be regarded as significant (on the order of -0.3% per year). If, on the other hand, it is found that it follows an ST process, then, given the high variance of the series in differences, the null hypothesis of a zero growth rate cannot be rejected.

As a first step towards determining which model is the most appropriate for each of the series to be analysed, both Augmented Dickey-Fuller (ADF) and non-parametric Phillips-Perron unit root tests are called for. The detailed results are presented in appendix 2. As the reader can see, after running the ADF tests, the null hypothesis for non-stationarity (existence of a unit root) cannot be discarded for seven of the eight indices and for 18 out of 24 products. When the Perron tests are run, the same thing occurs in the case of two indices and 14 products.[19]

These results are far from conclusive and, given their importance in determining the probability of obtaining a statistically significant trend, further work is called for in this area. In addition, the literature indicates that these two tests tend to lead to a false acceptance of the null hypothesis of a unit root, especially if the series has structural breaks.[20] What is more, if very small samples are being used and the shocks dissipate slowly, there may be very few independent observations of the process and, in that context, the estimation of DT models may generate more reliable parametrizations of the data (León and Soto, 1995b). In the case being discussed here, a number of authors have found proof of the existence of structural breaks or instability in the parameters.[21] In this instance the sample is finite and the speed at which the shocks dissipate is unknown.

For these reasons, an alternative method must be used in order to describe the persistence of a shock in the structure of the series. If its persistence is high, the series may be described as being non-stationary, since a shock will become a semi-permanent component of the series. On the other hand, if shocks dissipate rapidly, then the series would be stationary over time. This is tantamount to be non-parametric test for determining the existence of a unit root.

[18] It is important to emphasize that, whereas e_t is a random shock that does not affect the trend of the series (white noise), μ_t is a random variable whose presence induces stochastic behaviour in the trend.

[19] Based on the same data, Cuddington (1992) presents evidence that 12 of the 24 products can be modelled as non-stationary processes for the period 1900-1983. The results fit in with those presented in Table A.2, but with some important exceptions. For the period 1900-2000, coffee, lead and tin exhibit problems of non-stationarity that did not arise in Cuddington's results, which were also used by León and Soto (1995a). In addition, these authors find jute and rubber to be non-stationary based on the results of a Dickey-Fuller tests, but in this study they are regarded as being stationary, based on the 90% confidence interval yielded by the Perron test. If, however, we take a 95% confidence interval as a minimum significance requirement, as these authors did, then jute and rubber, as well as rice and the variable weightings index, would be non-stationary.

[20] León and Soto (1995a and 1995b) and Perron (1989).

[21] See, in particular, Cuddington et al. (2002). This question will be examined in greater depth in the following section.

In order to conduct such a test using the same approach as León and Soto (1995a and 1995b), recursive estimation procedures were employed to determine the ratio between the variance of innovation and the variance of the series. This estimator (known as Vk) makes it possible to see, from period to period (recursively), whether a shock changes the series' variability temporarily or permanently.[22] Its interpretation is explained in box 2. The Vk estimator also makes it possible to describe the response of the barter terms of trade based on a characteristic dissipation pattern (see section VIII).

Box 2
INTERPRETATION OF THE VK ESTIMATOR

The variance ratio can be used as a measurement of the significance of the permanent component as follows: If a series Yt follows a DT process, then no innovation has a permanent effect (i.e., the permanent component is null). Thus, in the long run, the variance of innovations and the Vk estimator will trend towards zero. If Yt is a random walk, then innovation is wholly captured by the permanent component, so the variance of innovations will tend to equal the variance of the series, and the variance ratio will be 1. In an intermediate process such as the ST process, Vk will be between 0 and 1.

The results of this estimation are presented in figure 2.[23] The thick lines depict the trend of the Vk estimator from period to period. The dotted lines trace its 95% confidence interval. If, throughout the period, Vk (or its confidence interval) trends towards 1, then the series is exhibiting a high level of persistence in the face of shocks and the series is therefore not stationary.[24] As may be seen from the Figure, this line of reasoning confirms the non-stationarity hypothesis for the prices of six products: cotton, aluminium, bananas, cocoa, silver and tobacco. The estimator does not provide conclusive evidence in the cases of tea and wool, and these series are therefore considered to be non-stationary, as has been indicated by earlier tests. In relation to the prices of the remaining products and the indices,[25] the null hypothesis for Vk trending towards 1 can be rejected and it can thus be concluded that they *do not exhibit non-stationarity*.[26]

Although the analysis conducted up to this point would lead us to believe that immediate ST estimation methodologies should be used for eight products (aluminium, bananas, cocoa, cotton, silver, tea, tobacco and wool) and that a DT model should be used for the remaining products and all of the indices, our basic hypothesis is that the deterioration in price indices occurred in a stepwise fashion, which would point to the presence of structural breaks in the series.[27] This question will be explored in greater depth in the following section.

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22 See Cochrane (1988); a detailed explanation of this procedure is given by León and Soto (1995b).

23 The program used to calculate this estimator was the program written by Paco Goerlich for RATS. This software is based on Cochrane (1988) (cochrane2.src) and is available at www.estima.com.

24 The way in which the estimator was constructed (León and Soto, 1995a and 1995b) causes the initial values to be close to 1, but what is important is its convergence towards or divergence from Vk=1.

25 The standard tests yielded biased results for coffee, beef, copper, tin, lead and the food products and metals indices.

26 If these results are compared with the findings of León and Soto (1995a), a number of differences arise. Whereas they found the trends for cocoa, silver and tea to be stationary during 1900-1992, the estimates made in the course of this study for 1900-2000 indicate that they are non-stationary.

27 Cuddington et al. (2002) contend that, regardless of whether a DT or ST specification is chosen, there is evidence that one or more breaks or instabilities in the parameters may be the problem.

Figure 2

RECURSIVE VK ESTIMATES OF PERSISTENCE

(Vertical axis: Vk; horizontal axis: time)

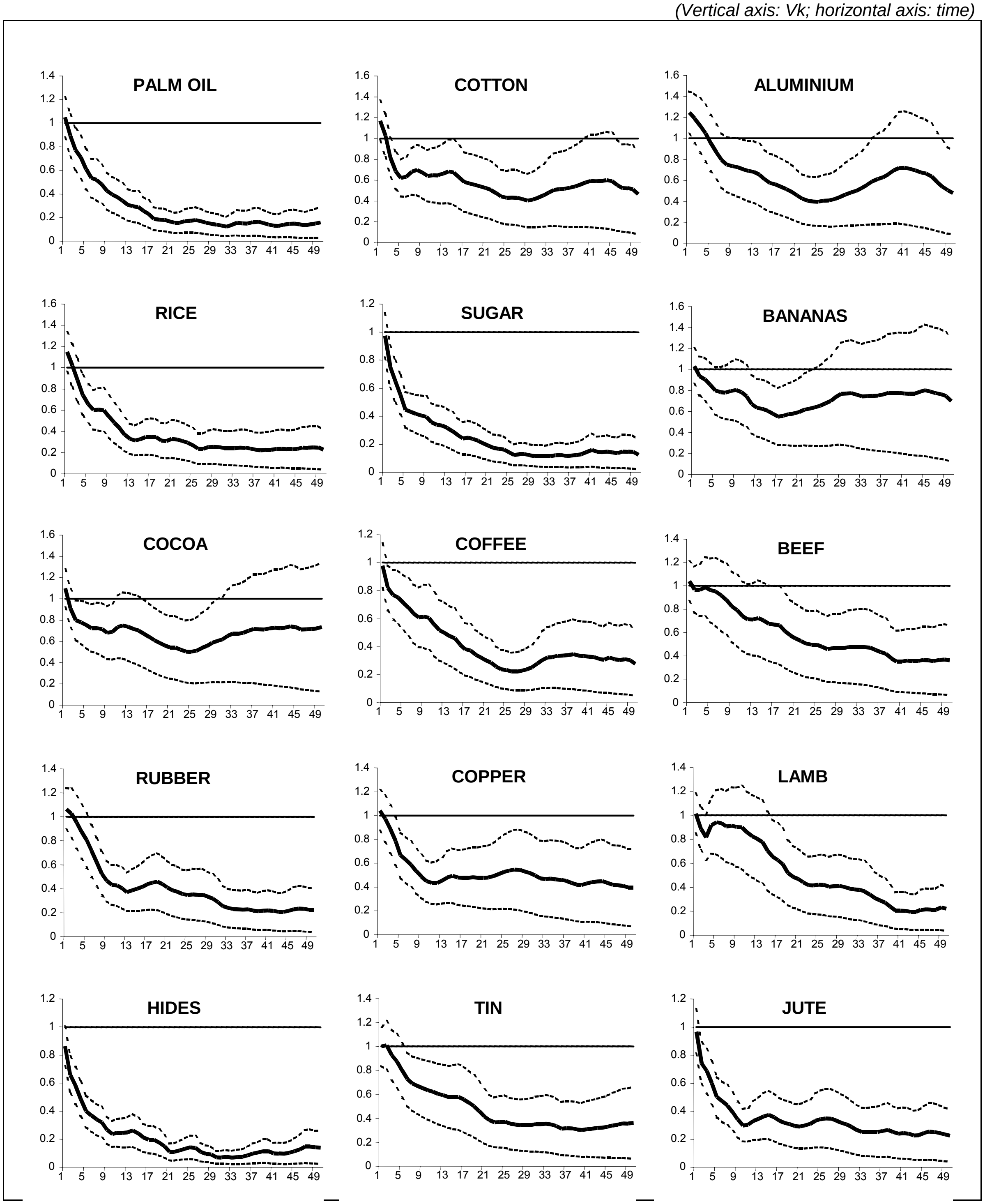
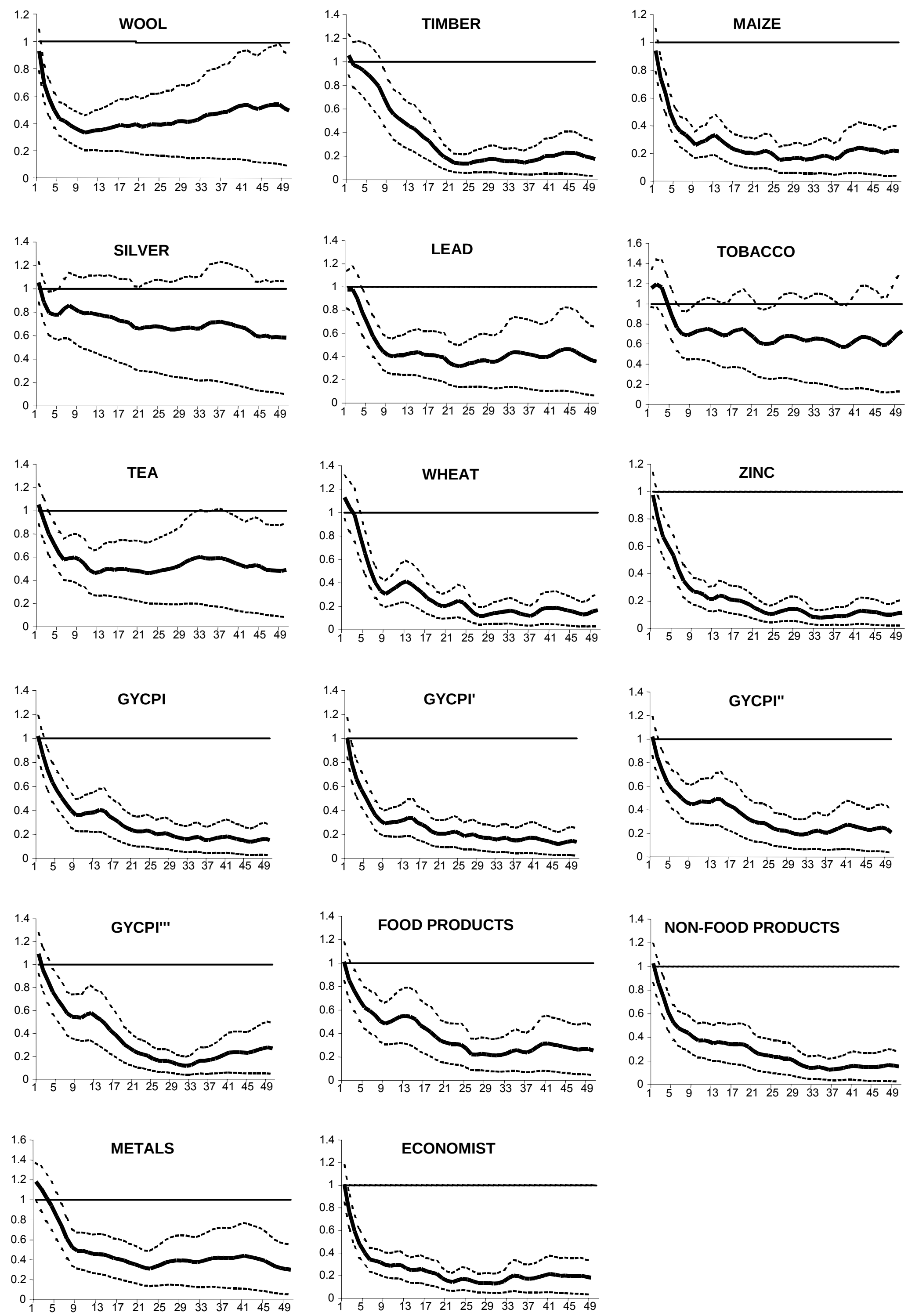


Figure 2 (continued)



Source: Calculations of the authors based on Grilli and Yang (1988), The Economist and ECLAC data.

IV. Structural breaks

The first step in this direction is to analyse the possible presence of structural breaks in the series that probably follow the deterministic-trend model.[28] Following Cuddington and others (2002), we calculate, first, the recursive residuals and the error bands for the hypothesis that the residuals come from the same distribution as those from the estimated model. We also show probabilities (p-values) for an N-step forecast test for each possible sample.[29] The results appear in appendix 3. As can be seen, for eight commodities (palm oil, sugar, beef, rubber, lamb, hides, timber and lead), these tests suggest the presence of a structural break around 1920. The same is true for all indices combined and the food subindex. In addition, five commodities and one index (palm oil, coffee, tin, lead, jute and GYCPI''') show evidence of a structural break around 1980. Lastly, rice, sugar, timber and the *Economist* index show evidence of a break around 1970.

To confirm the existence of these breaks, one of Perron's tests (1997)[30] may be used to conduct an endogenous search for the point where a structural break occurs. This is explained in box 3.

[28] An interesting overview of the work done in this area can be found in Cuddington et al. (2002).

[29] The null hypothesis is that forecast errors correspond to a model with no structural break. If the p-value is smaller than 0.01, then the null hypothesis can be rejected with a 99% confidence level.

[30] This procedure was written for the Regression Analysis of Time Series (RATS) software by G. Colletaz and F. Serranito of the Laboratoire d'économie d'Orléans, and is available at www.estima.com.

Box 3
PERRON'S TEST

This test selects a break point that minimizes the t-statistic for testing the null hypothesis of a unit root. The test may be estimated according to three possible models. The first allows only for a one-time change in the intercept, which occurs gradually. This is the *innovational outlier* (IO1) model. The second model allows for a change in both the intercept and the slope. This is the *innovational outlier with changing trend* (IO2) model. The third model allows for a change in the slope, with no break in the trend function. This is the *additive outlier* (AO) model.

Cuddington and others put forward four criticisms of this methodology and propose an alternative algorithm. The first criticism is that this is a test whose null hypothesis is the presence of a unit root, conditional on the possible presence of a structural break at an unknown date, and not a test for the presence of a structural break. The second is that it allows for only one structural break, whereas there is a priori no reason to believe that additional breaks may not be present. They also identify as a weakness the fact that the test allows for the structural break under the alternative hypothesis but not under the null hypothesis. Lastly, the test assumes that the type of structural break is known a priori. Notwithstanding these criticisms, the test has been used in this paper in view of its simplicity and the availability of the relevant algorithm.

Since we do not have a priori a structural form for each of the variables and we wish to avoid the last of the criticisms mentioned above, we shall apply the test to all three models for each series. The results for all the variables appear in appendix 4, table A.3. They are consistent with the results of the V^m estimator for six of the eight commodities that are non-stationary according to the estimator (cotton, bananas, cocoa, silver, tobacco and tea), since *it is not possible to reject the unit root hypothesis for these commodities* at a 95% confidence level.[31]

Eight commodities (coffee, beef, copper, lamb, tin, timber, maize and jute) and the GYCPI''' index *show no evidence of a structural break* according to this test, but they also show no evidence of a unit root according to the V^m estimator. Nonetheless, it must be borne in mind that, for most of these products,[32] this outcome is inconsistent with the results of the recursive residuals exercise (the case of these products will be re-examined in section VII).

Lastly, for eight commodities, the null hypothesis of a unit root can be rejected with a 95% confidence level with respect to the alternative hypothesis of a *structural break*. These commodities are palm oil, rice, sugar, rubber, hides, lead, wheat and zinc. The same is true of all the indices except the one that includes petroleum (GYCPI'''). These results are consistent with the recursive residuals analysis for all the commodities and indices except wheat, zinc and GYCPI'''. It is clear, however, that the presence of a structural break does not preclude the possible presence of one or more additional breaks. This possibility is considered in section VII.

[31] This is not the case for aluminium and wool.
[32] Except copper and maize.

Table 2 contains the results for these eight commodities and for the seven indices that show evidence of a structural break according to the Perron test. These results may at first appear confusing and, in terms of pinpointing a specific year for the structural break, inconclusive. One third of the changes take place between 1910 and 1930 and more than a third, between 1970 and 1990. While the selection of any time period is arbitrary, it is interesting to note that nine statistically significant changes can be detected in the period 1915-1925. Twelve more are found in the period 1973-1983 and eight more, in the period 1941-1951. Thus, two thirds of the breaks detected took place in these three time periods. Considering this evidence and, particularly, the fact that these periods coincide with the three historic turning points in the world economy (the two world wars and the end of the "golden age" of growth in the industrialized economies),[33] the deterministic and stochastic trend models will be estimated in the next section. The fact that breaks appear in more than one period will not be considered, however, until section VII.[34]

Table 2

TIMING OF STRUCTURAL BREAKS

(Vs. null hypothesis of unit root)

	Innovational outlier (intercept) IO1 1900-2000 RU	Innovational outlier (intercept and trend) IO2 1900-2000 RU	Additive outlier (trend) AO 1900-2000 RU
Hides	1950[b]	1916[b]	1905[b]
Lead	1978[a]	1945[b]	1973[b]
Palm oil	1917[b]	1983[c]	1991[b]
Rice	1988	1971[b]	1975[b]
Rubber	1915[b]	1934[b]	1926[a]
Sugar	1979[b]	1979	1983
Wheat	1941[c]	1928[b]	1910[c]
Zinc	1920[c]	1920[a]	1928[c]
GYCPI	1944[b]	1944[a]	1978[a]
GYCPI'	1948[b]	1971[b]	1929
GYCPI"	1983[b]	1970	1978[b]
Metals	1915[b]	1951[a]	1940[b]
Food	1983[b]	1984	1977[a]
Non-food	1948[b]	1938[b]	1929[a]
Economist	1915[c]	1915[c]	1920[c]

Source: Calculations of the authors.

[a] 90% significance.
[b] 95% significance.
[c] 99% significance.

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33 See Maddison (1995) and ECLAC (2002).

34 In particular, consideration will be given to those cases that show evidence of more than one structural break according to the recursive residuals test (palm oil, sugar, lead and the GYCPI" index).

V. Estimating the behaviour of real commodity prices

At this point, it would be useful to summarize the results obtained thus far:

- Trends in the prices of coffee, beef, copper, lamb, tin, timber, maize and jute and in the index that includes petroleum (GYCPI''') can be estimated according to a deterministic-trend model, using traditional econometric techniques;[35]
- Cotton, aluminium, bananas, cocoa, wool, silver, tea and tobacco should be estimated according to a stochastic-trend model;[36]
- The presence of a *structural break* should be considered in relation to the remaining indices and to palm oil, rice, sugar, rubber, hides, lead, wheat and zinc.

This section presents estimates of these models to determine whether there is a statistically significant trend in the series that follow a deterministic trend or a drift in those that follow a stochastic trend, and to identify the effect of structural breaks on the remaining series.

[35] The cases of coffee, beef, lamb, tin, timber and jute will be re-examined in section VII, as the two tests presented are not consistent and therefore do not preclude the possible presence of one or more structural breaks.

[36] An interesting observation is that five of these products (cotton, aluminium, cocoa, wool and tea) are among those that show the steepest falls, according to table 1. This will be taken into account in the subsequent analysis of the results of the stochastic-trend model estimate.

Section VII will cover, among other topics, the presence of more than one structural break and will accordingly re-estimate the models for all the variables that show evidence of at least one structural break in any of the tests.

Table 3 shows the results for the deterministic-trend model estimated by the ordinary least squares (OLS) method, adding ARMA so that the residuals will be white noise. From this, it may be concluded that coffee, copper and tin do not show a statistically significant deterministic trend. Beef, lamb and timber, on the other hand, have shown a positive trend of over 1% a year. Lastly, maize, jute and the aggregate index that includes petroleum (GYCPI''') show a constant declining trend. Of course, a model as simple as this one cannot completely capture the behaviour of the series.

Table 3

ESTIMATES OF VARIABLES THAT FOLLOW A DETERMINISTIC TREND

(Real variables reales in logarithms)

	C	β (%)	AR(1)	MA(1)	MA(2)	MA(4)	R^2
Beef	2.93[b]	**1.46**[a]	0.86[b]				0.88
Coffee	3.68[b]	0.21	0.81[b]				0.67
Copper	4.71[b]	-0.25	0.84[b]				0.72
Jute	5.40[b]	**-1.07**[a]	0.91[b]		-0.42[b]		0.72
Lamb	2.86[b]	**1.66**[b]	0.81[b]			0.31[b]	0.89
Maize	5.57[b]	**-1.29**[b]	0.70[b]				0.79
Timber	3.64[b]	**1.02**[b]	0.77[b]				0.87
Tin	3.64[b]	0.32	0.86[b]				0.76
GYCPI'''	5.08[b]	**-0.72**[b]	0.60[b]	0.44[b]			0.81

Source: Calculations of the authors.

[a] 95% significance.
[b] 99% significance.

Note: c: constant; β trend; AR(i): autorregresive dynamics of order i; MA(i): moving average dynamics of order i; R^2: coefficient of determination.

Table 4 shows the results for the eight variables that manifest non-stationarity. In view of their wide variability, it is unsurprising that none of the drifts is statistically significant.[37] Their sign, however, is *negative* for all of these commodities except tobacco. If these results are compared to those in table 1, it becomes apparent that the five products that show a drift[38] of close to -1% accumulated a decline of nearly 60% between 1900-1904 and 1996-2000. For these prices, then, the frequency of negative shocks has clearly outweighed the effects of positive ones, with the result that they have deteriorated significantly.

[37] Cuddington and others reach the same conclusion using the stochastic-trend model for the GYCPI series.
[38] Not statistically significant.

Table 4
STOCHASTIC-TREND MODEL ESTIMATES
(Differences in real variables in logarithms)

	C (%)	AR(2)	MA(1)	MA(2)	MA(3)	MA(4)	R^2
Aluminium	-1.10		0.28[c]	-0.20[b]			0.13
Bananas	-0.01			-0.21[b]			0.03
Cocoa	-1.20	-0.33[c]					0.11
Cotton	-1.13			-0.28[c]	-0.19[a]		0.14
Silver	-0.26			-0.26[b]			0.07
Tea	-0.82	-0.24[b]					0.06
Tobacco	0.77					-0.27[c]	0.08
Wool	-1.40			-0.41[c]			0.13

Source: Calculations of the authors.

[a] 90% significance.
[b] 95% significance.
[c] 99% significance.

Note: c: constant; AR(i): autorregresive dynamics of order i; MA(i): moving average dynamics of order i; R^2: coefficient of determination.

Lastly, table 5 shows the results obtained by estimating the three models considered by Perron (see box 3). This Table is organized on the assumption that a structural break occurs at time Tb. Accordingly, it shows the values of both the intercept and the trend, both before and after the structural break occurs.[39]

According to these results, the relative price trend for palm oil, rice, sugar,[40] rubber, hides, lead and wheat has been systematically *negative*.[41] The trend for zinc has been systematically *positive*, but exhibited a steep fall in 1920. The behaviour of the β value indicates that the trend has been *constant and negative* only in the case of sugar. In the other cases, β obscures the fact that, for some commodities, the trend was not significant prior to the structural break (as in the cases of rice, hides, lead and wheat), whereas for others it lost significance after the break (as in the cases of palm oil and rubber (IO2)).

[39] For example, before 1917, the intercept for palm oil was 4.96 and later shifted to 5.09. The trend, meanwhile, was -0.72%, taking this break into account. The trend changed after 1991 from -0.46% to -3.4% per year, while the intercept remained at 4.90.
[40] This variable exhibited a sharp decline in 1979.
[41] Only the segments in which the trend is statistically significant have been considered.

Table 5

ESTIMATES OF THE DETERMINISTIC-TREND MODEL WITH STRUCTURAL BREAKS, FOR INDIVIDUAL PRODUCTS

	Intercept			Trend					
	[1900, Tb]	[Tb, 2000]	C	[1900, Tb]	[Tb, 2000]	(%)	AR(1)	MA(1)	R^2
Hides									
IO1, 1950	5.21[c]	5.14[c]				-1.09[c]	0.57[c]		0.72
IO2, 1916			5.07[c]	0.22	-1.24[c]		0.56[c]		0.73
AO, 1905			5.19[c]	-4.23	-1.21[c]		0.56[c]		0.72
Lead									
IO1, 1978	4.75[c]	4.84[c]				-0.91[a]	0.86[c]		0.76
IO2, 1945			4.50[c]	-0.10	-1.21[c]		0.79[c]		0.77
AO, 1973			4.42[c]	0.05	-2.99[c]		0.67[c]		0.79
Palm oil									
IO1, 1917	4.96[c]	5.09[c]				-0.72[c]	0.52[c]	0.37[c]	0.66
AO, 1991			4.90[c]	-0.46[b]	-3.40		0.55[c]	0.37[c]	0.65
Rice									
IO2, 1971			4.98[c]	-0.08	-4.02[c]		0.47[c]	0.54[c]	0.86
AO, 1975			5.05[c]	-0.33[a]	-5.27[c]		0.47[c]	0.51[c]	0.87
Rubber									
IO1, 1915	6.32[c]	6.83[c]				-2.40[c]	0.74[c]		0.92
IO2, 1934			5.87[c]	-3.00[c]	-1.46		0.88[c]		0.91
AO, 1926		6.24[c]		-3.33[c]	-4.29[c]		0.87[c]	0.24[b]	0.92
Sugar									
IO1, 1979	5.25[c]	4.99[c]				-0.72[b]	0.38[c]	0.46[c]	0.64
Wheat[d]									
IO1, 1941	5.37[c]	5.40[c]				-0.92[c]	0.38[c]	0.62[c]	0.81
IO2, 1928			5.20[c]	-0.01	-1.02[c]		0.35[c]	0.65[c]	0.81
AO, 1910			5.31[c]	-0.95	-0.91[c]		0.34[c]	0.66[c]	0.81
Zinc									
IO1, 1920	4.72[c]	4.35[c]				0.39[c]		0.55[c]	0.46
IO2, 1920			4.46[c]	2.66[c]	0.32[c]			0.53[c]	0.47

Source: Calculations of the authors.

[a] 90% significance. [b] 95% significance. [c] 99% significance. [d] The error structure of the equations reveals a misspecification, since ARMAs of more than 2 are needed to correct the autocorrelation, and the AR(1) coefficient is greater than 1.

Note: The definition of the IO1, IO2 and AO models is given in box 3. β: trend; AR(i): autorregresive dynamics of order i; MA(i): moving average dynamics of order i; R^2: coefficient of determination.

Table 6 presents a similar analysis for the non-petroleum aggregate price indices. Clearly, all of them show a systematic negative trend. However, for GYCPI, GYCPI” and the food sub-index, this trend was not significant prior to the structural break. Another noteworthy finding is the proportionate increase in the deterioration of GYCPI’, the metals sub-index and the non-food sub-index after each structural break.

As little long-term information is available on productivity, transport costs and changes in product quality comparable to the price series examined above, it is difficult to include these variables in the statistical exercises. In any event, existing productivity series for the OECD countries show a break in the trend of relative labour productivity in agriculture and manufacturing in the 1950s: while manufacturing productivity had risen faster than agricultural productivity up to that time, the opposite has been the case since the 1950s (Bairoch, 1989; Maddison, 1991). This structural break is not, however, reflected in the foregoing statistical results. Furthermore, the long-term lead gained by agricultural productivity, as revealed by these data, would only explain a relatively marginal decline in agricultural terms of trade (of about 0.2% a year). There are no comparable series for the developing world, and even if such series were available, they would be distorted by changes in the rural underemployment that characterized the developing countries throughout the twentieth century.

Table 6

ESTIMATES OF THE DETERMINISTIC-TREND MODEL WITH STRUCTURAL BREAKS, PRICE INDICES

	Intercept			Trend					
	[1900, Tb]	[Tb, 2000]	C	[1900, Tb]	[Tb, 2000]	β (%)	AR(1)	MA(1)	R^2
GYCPI									
IO1, 1944	5.05[c]	5.11[c]				-0.81[c]	0.71[c]		0.83
IO2, 1944			4.85[c]	-0.17	-1.02[c]		0.68[c]	0.25[a]	0.83
AO, 1978			4.81[c]	-0.19	-2.94[c]		0.70[c]	0.22[a]	0.83
GYCPI'									
IO1, 1948	5.17[c]	5.32[c]				-1.01[c]	0.41[c]	0.41[c]	0.86
IO2, 1971			4.94[c]	-0.34[b]	-2.54[c]		0.57[c]	0.34[c]	0.85
GYCPI"									
IO1, 1983	5.16[c]	4.88[c]				-0.72[c]	0.58[c]	0.30[b]	0.89
AO, 1978			4.94[c]	-0.20	-4.42[c]		0.67[c]	0.27[b]	0.88
Metals									
IO1, 1915	4.79[c]	5.14[c]				-0.80[b]	0.85[c]	0.30[c]	0.86
IO2, 1951			4.85[c]	-0.39[a]	-0.94[b]		0.79[c]	0.36[c]	0.85
AO, 1940			4.89[c]	-0.56[a]	-0.86[b]		0.80[c]	0.42[c]	0.85
Food									
IO1, 1983	4.89[c]	4.54[c]				-0.46[b]	0.72[c]		0.83
AO, 1977			4.76[c]	-0.14	-3.94[c]		0.51[c]	0.35[b]	0.84
Non-food									
IO1, 1948	5.20[c]	5.32[c]				-1.10[c]	0.52[c]	0.32[b]	0.86
IO2, 1938			5.01[c]	-0.60[b]	-1.30[c]		0.62[c]	0.35[c]	0.86
AO, 1929			4.94[c]	0.34	-0.95[c]		0.64[c]	0.31[b]	0.86
Economist									
IO1, 1915	6.75[c]	7.04[c]				-1.33[c]		0.67[c]	0.91
IO2, 1915			6.82[c]	-1.58[b]	-1.26[c]			0.72[c]	0.90
AO, 1920			6.75[c]	-0.13	-1.27[c]		0.28[a]	0.53[c]	0.91

Source: Calculations of the authors.
[a] 90% significance. [b] 95% significance. [c] 99% significance.

Note: The definition of the IO1, IO2 and AO models is given in box 3. β: trend; AR(i): autorregresive dynamics of order i; MA(i): moving average dynamics of order i; R^2: coefficient of determination.

Series on productivity per hectare for seven agricultural products, estimated by Scandizzo and Diakosawas (1987) and updated with information from the Food and Agriculture Organization of the United Nations (FAO), indicate annual productivity improvements of around 1% throughout the twentieth century, with the notable exception of coffee, for which productivity increased by just 0.2% a year between 1910-1914 and 1995-1999. The rate of productivity growth sped up between the 1960s and the 1980s for three products that were affected by the "green revolution" (rice, maize and wheat). The inclusion of these productivity series in the statistical exercises does not, however, change the conclusions to be drawn concerning long-term real price trends; in fact, their impact on real prices seems to have been only partial and not always statistically significant. There are perhaps just two exceptions to this rule. First, the absence of any long-term adverse trend in coffee prices may be related to its slower productivity growth. Second, the structural break in real prices for rice in the early 1970s may be associated with the green revolution (though maize and wheat did not show any such breaks).

In sum, coffee, copper and tin show *no statistically significant deterministic trend*. Likewise, cotton, aluminium, bananas, cocoa, wool, silver, tobacco and tea exhibit a drift that is not statistically significant. All of them, however, except silver and tobacco, have suffered a *cumulative deterioration of nearly 60%*. On the other hand, beef, lamb and timber have shown a *positive* deterministic trend. Zinc has also followed a systematically positive trend. Lastly, maize, jute and the aggregate index GYCPI''' have followed a *constant trend towards deterioration*, while the relative price trend for palm oil, rice, sugar, rubber, hides, lead and wheat and of all the other aggregate indices has been systematically *negative*. Therefore, 4 commodities show a positive trend; 11 show no trend or drift, though 5 of them experienced a cumulative decline of nearly 60%, indicating the strong predominance of negative shocks; and 9 show a negative trend. In the aggregate, commodities with negative trends or shocks prevail and, thus, all indices tend to deteriorate.

In general, given the scarcity of available information, no firm conclusions can be drawn about the impact of relative agricultural productivity on the long-term trend —or the breaks in the trend— of the terms of trade for agricultural products in the twentieth century; even less can be inferred about the impact on their factorial terms of trade. As pointed out in section III, the literature on the convergence or divergence of real wages and per capital income is more relevant for assessing the validity of what is referred to here as the second Prebisch-Singer hypothesis.

VI. Multiple structural breaks

The estimates of models in the preceding section, using Perron's methodology (1997), reveal that, in the case of eight commodities and seven of the eight indices, there was at least one structural break in price trends in the course of the twentieth century. They also confirm that these breaks have tended to erode the commodity terms of trade. This tends to substantiate our basic hypothesis that the deterioration of those prices took place in a stepwise fashion. Unfortunately, the methodology used so far allows for only a one-time structural break.[42] This means that there is still room for another possible break to "hide" behind the statistical estimates. Indeed, this is precisely what is suggested by the results of the recursive residuals analysis and the fact that the breaks are located around more than one period (see section V).

While the fact cannot be established with strict econometric rigour, the foregoing results and economic history itself imply that the biggest breaks were concentrated around 1920 and 1980. This suggests that they represent the delayed effects of the sharp slowdowns in the world economy after the First World War and after the first oil shock of the 1970s, which marked the end of the "golden age" of the industrialized economies (Madisson, 1995). More precisely, the following econometric analysis assumes that the breaks took place around 1921 and 1979, coinciding, respectively, with the severe international crisis that followed the First World War, whose effect on raw materials prices is well known, and with the monetary shock generated by the actions of the United States economic authorities to

42 Cuddington et al. present the results of a model that accommodates more than one structural break for the GYCPI series. However, that model could not be applied on a large scale to all the series included in this analysis owing to its econometric complexity.

curb inflation. Statistical exercises were also performed to determine whether there was a structural break in the series at the end of the Second World War or shortly thereafter (around the time of the Korean War). Since the results did not point to a significant statistical break in that time period, they have not been reported in this paper.

Accordingly, these results and historical facts warrant the conduct of a final econometric exercise involving the re-estimation of all the models[43] except those that show a stochastic trend, assuming that structural breaks took place in 1921 and 1979. Table 7 contains the corresponding estimates, and figure 3 shows the results for the indices, excluding the ARMA dynamic of the residuals to illustrate more clearly the breaks and the deviations of prices from the estimated trends.

These exercises show that food became considerably more expensive up until the First World War, whereas metals lost value. In terms of specific commodities, only a few (palm oil, hides, timber and maize, in particular) followed a rising trend, and only one (rubber) followed the opposite pattern.

In 1921 all the Grilli-Yang aggregate indices experienced a *strong and sudden decline* (of between 44% and 52%, depending on the aggregate index used), from which they failed to recover in subsequent decades. In terms of product groups, metals were the only exception to this rule. The decline is statistically significant, though of varying magnitude, for 11 out of 16 commodities. Interestingly, this drop was followed by a long period (1922-1979) in which aggregate price indices followed *no* statistically significant trend. This occurred because the different prices moved in opposite directions.

Finally, in contrast to what had happened in 1921, in 1979 there was no sudden drop in prices, but rather *a break in the price trend*, which became strongly negative from then on (with declines of 2% to 3% a year for the various sub-indices). This trend was very pronounced for food and less so for metals; moreover, it was negative for 14 of the 16 commodities included in table 7, though it was statistically significant for only 9. A closer analysis might indicate that the decline was concentrated in the 1980s (see Maizels, 1992, for an analysis of changes in raw materials prices in that decade), in which case this phenomenon would be more similar to what took place in 1921, though it was more gradual over time.

Lastly, it should be pointed out that the behaviour of the *Economist* series also involves sudden shifts, though it differs from that of the Grilli-Yang indices. In particular, the 1921 adjustment is smaller (20%), but the series shows a strong, statistically significant negative trend in the period 1922-1979 (1.2% a year), which picked up speed after 1979. Thus, as illustrated in figure 3, this series follows a much more secular trend towards deterioration starting in the 1920s.

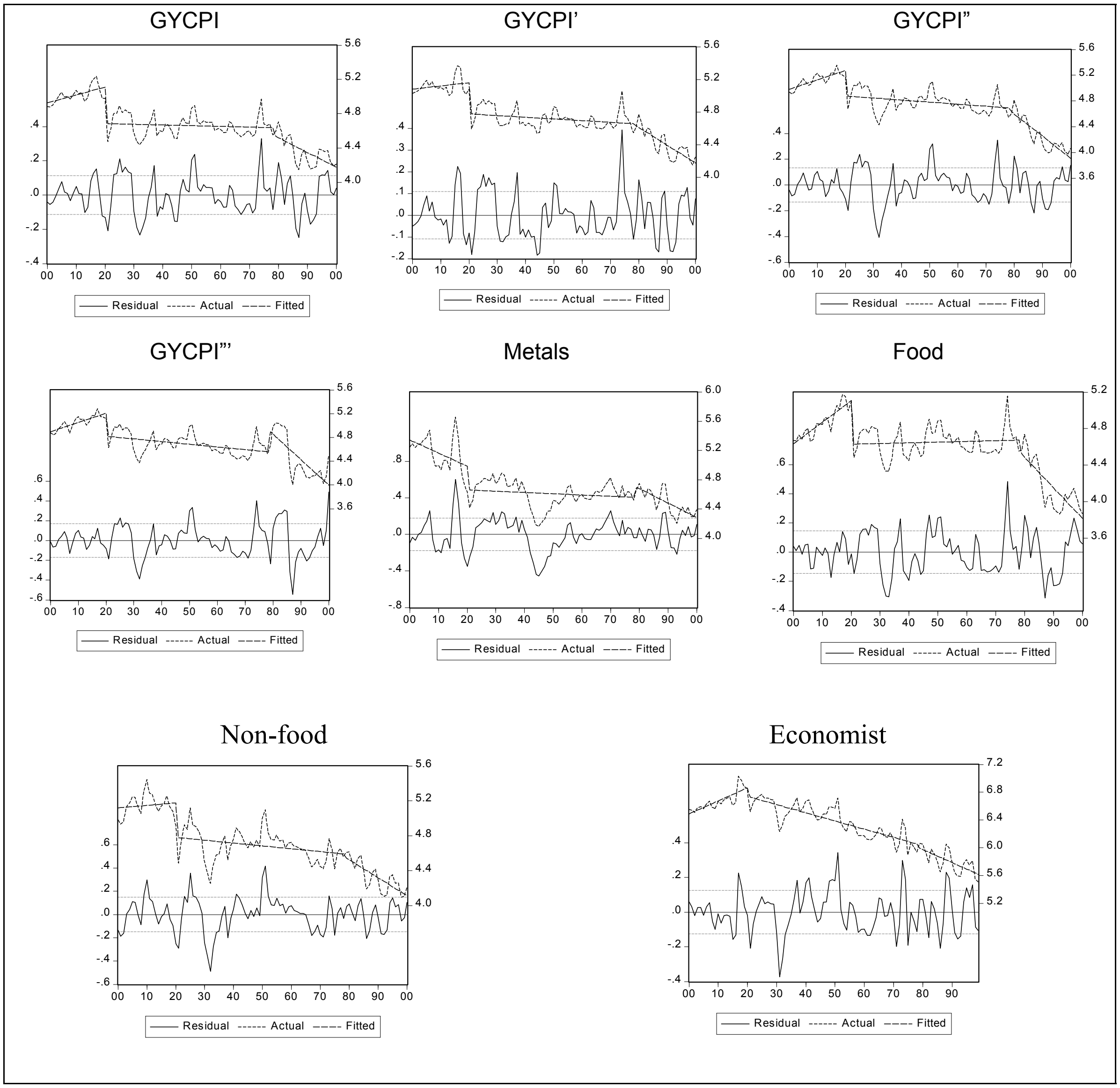
The statistical exercises therefore seem to indicate that the reduction in real commodity prices throughout the twentieth century resulted from two major structural breaks that took place around 1921 and 1979. The first of these breaks took the form of a sudden, one-time drop in prices, and the second took the form of a shift in the trend of commodity prices.

To complete this overview, the following section contains a brief analysis of the series' speed of mean reversion in response to short-term shocks. A slow speed of mean reversion would imply that short-term shocks have a long-lasting effect on economic performance.

[43] In the preceding sections justifications for re-estimating all the stationary price series except copper and maize were presented. These two products have been included to round out the analysis by covering all stationary series. Non-stationary variables cannot be estimated using traditional methods because, as shown earlier, these methods may generate spurious results.

Figure 3

ESTIMATION OF AGGREGATE PRICE INDICES WITH MULTIPLE STRUCTURAL BREAKS



Source: Calculations of the authors.

Notes: **GYCPI**: Total index, weighted by the share of total exports represented by each product in 1977-1979; three subindices are also derived: *food products, non-food products* and *metals*.

GYCPI': Total index, weighted by the developing countries' share of commodity exports in 1981. (The original index used weightings for 1977-1979; since these weightings were unavailable, weightings for 1981 were substituted.)

GYCPI'': Total index, weighted by the share of world exports represented by commodities during the year in question.

GYCPI''': The same as GYCPI'' except that it also includes oil prices.

Table 7

ESTIMATION WITH STRUCTURAL BREAKS IN 1921 AND 1979

	c	t021	c21	t2279	c79	t8000	AR(1)	AR(2)	AR(4)	MA(1)	MA(2)	MA(4)
Beef	3.03	0.72	-0.18	2.49[c]	0.32	-4.28[b]	0.80					
Coffee	3.14	3.74	-0.41	1.39[b]	-0.11	-4.11[b]	0.73					
Copper	5.03	-1.38	-0.42[c]	0.85[c]	-0.10	-2.22[b]				0.81	0.33	
Hides[d]	4.82	3.29[c]	-0.63[c]	-1.10[c]	0.37[b]	-3.58[c]				0.52		
Jute	4.57	2.66[a]	-0.28	0.19	-0.32[a]	-2.90[b]	0.31			0.51		
Lamb	2.66	3.23	-0.16	1.84[b]	0.34	-1.46	0.79					0.36
Lead	4.35	1.11	-0.28[a]	0.22	0.14	-5.60[c]	0.76					
Maize	4.90	3.30[c]	-0.44[c]	-0.54[c]	-0.20	-3.73				0.55		
Palm oil	4.67	3.20[b]	-0.53[c]	-0.41	-0.27[a]	1.12				0.82	0.27	
Rice	5.28	-1.86	0.12	-0.42[a]	-0.31[b]	-3.13[c]				0.91	0.26	
Rubber	7.38	-6.62[c]	-0.67[b]	-1.06[c]	-0.33[a]	-2.00	1.62	-0.74		-1.00		
Sugar	5.20	2.58	-1.04[c]	0.34	-0.33	-2.01	-0.48			1.39	0.63	
Timber	3.16	5.22[c]	-0.41[c]	1.08[c]	-0.00	0.76	0.70					
Tin	3.50	1.17	-0.36[a]	1.64[c]	0.09	-5.01[c]	0.67					
Wheat	5.08	1.72[a]	-0.31[b]	-0.66[c]	-0.03	-1.78[b]	0.35		-0.26	0.52		
Zinc	4.65	0.81	-0.42[c]	0.59[b]	-0.03	-0.32				0.69	0.21	
Indices												
GYCPI	4.91	1.24	-0.49	-0.08	-0.06	-1.94[c]				0.82	0.29	
GYCPI'	5.06	0.67	-0.44	-0.18	-0.02	-2.15[c]				0.80	0.25	
GYCPI''	4.95	1.84	-0.48	-0.29[a]	-0.06	-3.35[c]				0.85	0.29	
GYCPI'''	4.86	2.10	-0.52	-0.30	0.17	-3.09[c]				0.98	0.38	
Metals	5.46	-2.82	-0.19	-0.21	0.12	-1.66[a]	0.59			0.41		
Food	4.57	3.05	-0.60	0.09	-0.11	-3.61[c]	0.37			0.51		
Non-food	5.11	0.45	-0.44	-0.33[a]	-0.01	-2.19[c]				0.82	0.30	
Economist	6.47	2.12	-0.20	-1.17[c]	-0.02	-2.06[c]				0.67		

Source: Calculations of the authors.

[a] 90% significance.
[b] 95% significance.
[c] 99% significance.
[d] Does not converge.

Note: c: constant; AR(i): autorregresive dynamics of order i; MA(i): moving average dynamics of order i.

VII. Variability and short-term shocks

The V^m estimator, which was used in section IV to determine the long-term persistence of innovations, is also useful for analysing the series' reaction to short- and medium-term shocks without resorting to methodologies based on parametrizations that give too much weight to short-term movements. The speed with which the estimator tends towards zero shows how a shock is dissipated. Following the methodology of León and Soto (1995b), table 8 illustrates the behaviour of this estimator for variables that do not present problems of non-stationarity.

This analysis reveals that nine commodities show a significant mean reversion process in the first five years after a shock occurs. In six cases (sugar, hides, jute, wool, maize and zinc), the shock has dissipated by 40% within the first four years. In three cases (palm oil, rice and coffee), this reduction amounts to 25%. When the same parameters are applied to the indices, all of them except the index that includes petroleum and the metals price index show a high speed of mean reversion. After this initial reversion, the process continues at a slower pace, so that after more than 25 years only nine commodities have returned to long-term equilibrium (V^m<0.26).

Viewed from a macroeconomic standpoint, these results show that, despite the relative speed of mean reversion, the effects of a shock last for more than a year, and therefore have an impact in both the short and medium terms. This indicates that stabilization funds are a viable option, but the reference prices used by such funds must change according to market prices to prevent large-scale fiscal losses associated with their management.

Table 8

ESTIMATION OF THE MEAN REVERSION PROCESS

(Value of the V^m statistic)

	Years							
	1	2	3	4	5	10	15	28
Beef	1.04	0.97	0.97	0.99	0.96	0.79	0.70	0.47
Coffee	0.98	0.82	0.77	0.75	0.71	0.61	0.46	0.23
Copper	1.04	0.97	0.89	0.78	0.67	0.45	0.49	0.54
Hides	0.86	0.66	0.58	0.49	0.40	0.24	0.24	0.10
Jute	0.97	0.74	0.69	0.60	0.50	0.30	0.37	0.33
Lamb	1.01	0.89	0.82	0.92	0.94	0.90	0.70	0.41
Lead	0.98	0.98	0.89	0.77	0.67	0.40	0.44	0.37
Maize	0.94	0.74	0.64	0.50	0.40	0.28	0.27	0.16
Palm oil	1.05	0.90	0.77	0.71	0.61	0.39	0.29	0.16
Rice	1.15	1.02	0.88	0.74	0.66	0.48	0.33	0.25
Rubber	1.07	1.03	0.95	0.87	0.80	0.43	0.40	0.35
Sugar	0.98	0.75	0.65	0.55	0.45	0.36	0.27	0.13
Timber	1.06	0.97	0.96	0.93	0.89	0.56	0.40	0.17
Tin	1.00	1.01	0.93	0.88	0.80	0.63	0.58	0.34
Wheat	1.13	1.05	0.98	0.81	0.65	0.33	0.35	0.12
Wool	0.93	0.69	0.58	0.49	0.43	0.33	0.37	0.41
Zinc	0.98	0.80	0.67	0.60	0.53	0.26	0.23	0.14
GYCPI	1.02	0.87	0.74	0.63	0.57	0.36	0.40	0.21
GYCPI'	1.01	0.79	0.66	0.58	0.50	0.30	0.34	0.20
GYCPI"	1.02	0.85	0.73	0.64	0.58	0.46	0.49	0.22
GYCPI"'	1.10	0.95	0.86	0.76	0.70	0.54	0.49	0.16
Metals	1.18	1.11	1.02	0.94	0.84	0.49	0.44	0.39
Food	1.01	0.85	0.77	0.68	0.62	0.50	0.52	0.23
Non-food	1.03	0.86	0.74	0.61	0.52	0.37	0.36	0.22
Economist	1.01	0.76	0.59	0.48	0.40	0.29	0.25	0.13

Source: Calculations of the authors.

VIII. Conclusions

The econometric results presented in this paper do not provide evidence of a secular or continuous trend towards the erosion of the terms of trade. It is nevertheless a fact that relative raw materials prices deteriorated markedly in the course of the twentieth century. Various tests confirm that there has been a decline, whether escalated or continuous, in the prices of nine commodities and in all the indices. Meanwhile, eight commodities reveal the presence of a unit root and high volatility; it is therefore not surprising that their drift, though negative for all but one of the products, is not statistically significant. At the same time, however, the cumulative decline for five of them amounts to nearly 60%, indicating that negative shocks far outweighed positive ones. Lastly, four products follow a rising trend, and three others show no significant deterministic trend.

Since all the non-petroleum indices show evidence of structural breaks, both the information generated from the tests performed and the evidence of historical facts can be used to affirm that the first abrupt downward shift seems to have taken place around 1920 and was related to the major global economic changes produced by the First World War. The second structural break seems to have occurred around 1980, in the wake of the world economic slowdown that began in 1973. Econometric analysis confirms the presence of these breaks in different ways. Whereas the first case involved a one-time adjustment, of significant magnitude, in commodity prices, the second case involved an adverse shift in the trend of global prices. Prices followed a more positive trend before the First World War, and there is no clear evidence of a significant trend in commodity prices between the 1920s and the 1970s.

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Appendices

Appendix 1

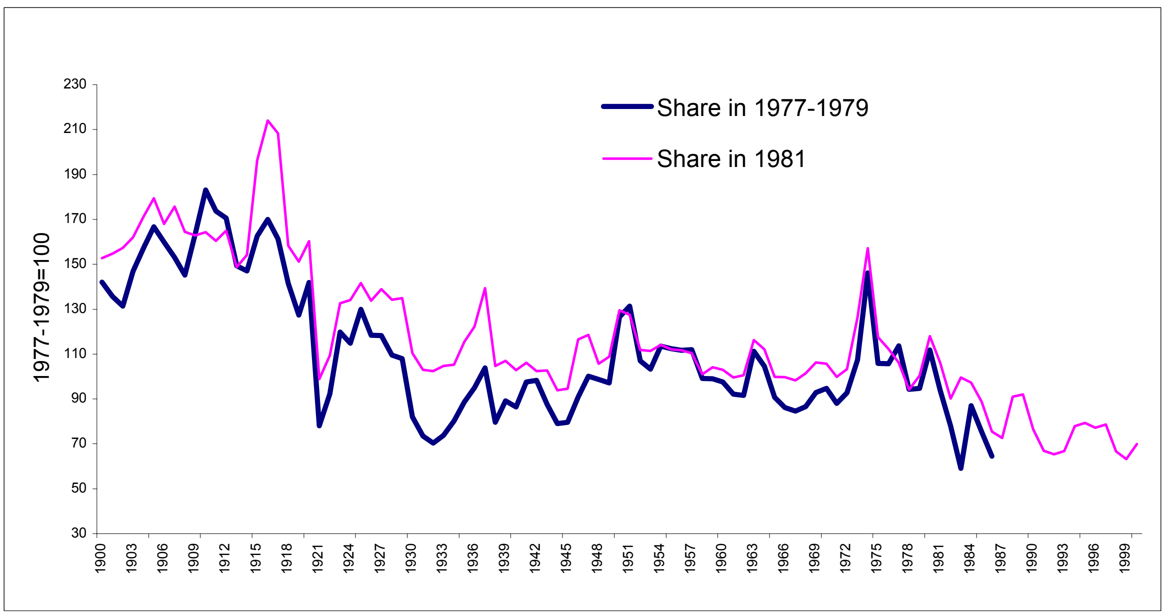
Methodology used to update price indices

John Cuddington generously provided data on the prices of the 24 different products used by Grilli and Yang (1988). These data were then updated for the period from 1987 on by ECLAC based on statistics compiled by UNCTAD, the International Monetary Fund, the World Bank and OPEC. These series were then used to update the price indices (four aggregate indices and three subindices) developed by Grilli and Yang (1988).

The first of these (GYCPI) indices used weightings for the share of world exports represented by each commodity in 1977-1979. These weightings, which were published by Cuddington and Wei (1998), were used to update the index to 2000. The second index (GYCPI') was weighted by the developing countries' share of commodity exports in 1977-1979. The original weightings are unavailable, so weightings were developed for 1981 based on COMTRADE data; the index was then recomputed for the entire century. Figure A1 shows the original and the new series, in real terms. The differences between the two are concentrated in the years prior to 1950. Nonetheless, the trends are similar except during the years of the First World War, when the new weightings amplify the upswing in the price index.

Figure A.1

REAL PRICE INDEX, WEIGHTED BY DEVELOPING COUNTRIES' SHARE OF COMMODITY EXPORTS



Source: Grilli and Yang (1988) and calculations based on United Nations data.

The other two of Grilli and Yang's aggregate indices (GYCPI'' and GYCPI''') have variable weightings based on the share of world exports represented by commodities in different years. The difference between the two is that the second index includes petroleum. Variable weightings for the current year were calculated by ECLAC based on COMTRADE data and were then used to update these indices. The results are shown in table A1.

Table A.1

UPDATING OF GRILLI AND YANG'S PRICE INDICES

(1977-1979 = 100)

	GYCPI	GYCPI'	GYCPI''	GYCPI'''	GYCPI subindices		
					Food	Non-food	Metals
1986	88.36	98.42	90.79	93.76	75.78	102.58	134.42
1987	93.61	107.56	92.56	81.28	90.54	124.55	180.99
1988	118.92	142.86	117.15	114.38	91.50	128.97	178.08
1989	123.29	143.19	116.13	120.54	86.67	125.89	157.21
1990	120.81	131.66	108.60	132.39	83.97	106.87	128.75
1991	109.78	115.15	96.77	118.55	84.44	107.25	129.28
1992	113.93	115.63	95.79	110.09	82.80	102.75	111.69
1993	109.25	111.43	91.96	102.60	102.04	131.04	133.96
1994	138.01	133.97	108.45	109.15	105.53	146.73	157.42
1995	149.71	149.57	117.79	118.84	108.12	128.56	141.00
1996	141.99	140.56	108.81	121.87	107.14	119.68	141.69
1997	133.75	133.36	107.57	118.27	93.35	98.84	120.37
1998	112.35	109.70	90.82	92.25	81.77	98.06	118.00
1999	107.01	102.46	85.61	101.53	75.03	105.75	126.19
2000	105.57	109.50	90.87	138.72	73.10	103.67	120.80

Source: Calculations of the authors.

GYCPI: Dollar-denominated price index for 24 tradable non-fuel commodities, weighted by the share of total exports represented by each product in 1977-1979.

GYCPI': Dollar-denominated price index for 24 tradable non-fuel commodities, weighted by the developing countries' share of commodity exports in 1981.

GYCPI'': Dollar-denominated price index for 24 tradable non-fuel commodities, weighted by the share of world exports represented by commodities during the year in question.

GYCPI''': The same as GYCPI'' except that it also includes oil prices.

Appendix 2

Table A.2

UNIT ROOT TESTS FOR THE LOGARITHM OF THE SERIES IN REAL TERMS

	Augmented Dickey-Fuller statistic	Phillips-Perron statistic
Aluminium	-2.21	-2.40
Bananas	-2.04	-2.59
Beef	-2.84	-2.78
Cocoa	-2.21	-2.46
Coffee	-2.55	-3.05
Copper	-2.23	-2.98
Cotton	-1.72	-2.38
Hides	-3.71 [b]	-5.10 [c]
Jute	-2.31	-3.24 [a]
Lamb	-3.50 [b]	-3.06
Lead	-1.94	-2.72
Maize	-2.49	-4.21 [c]
Palm oil	-3.99 [b]	-4.22
Rice	-2.41	-3.16
Rubber	-3.03	-3.34
Silver	-1.98	-2.31
Sugar	-3.09	-4.49
Tea	-1.80	-2.39
Timber	-3.98 [b]	-3.80
Tin	-2.33	-2.70
Tobacco	-1.05	-1.90
Wheat	-3.90 [b]	-4.37
Wool	-2.04	-2.80
Zinc	-4.09 [c]	-4.84
GYCPI	-2.82	-3.86
GYCPI'	-2.83	-4.09
GYCPI''	-2.25	-3.29
GYCPI'''	-3.06	-3.67
Food products	-2.12	-2.97
Non-food products	-2.67	-3.91
Metals	-2.94	-3.13
Economist	-3.64 [b]	-3.89

Source: Calculations of the authors.
[a] 99% significance.
[b] 95% significance.
[c] 90% significance.
Four lags are used for the ADF.
Sign: Significance in the event that the null hypothesis is rejected.
Product names in boldface type correspond to those series for which the null hypothesis could not be rejected on the basis of either of the two tests.

Appendix 3

Figure A.2

RECURSIVE RESIDUALS TESTS OF THE EXISTENCE OF STRUCTURAL BREAKS

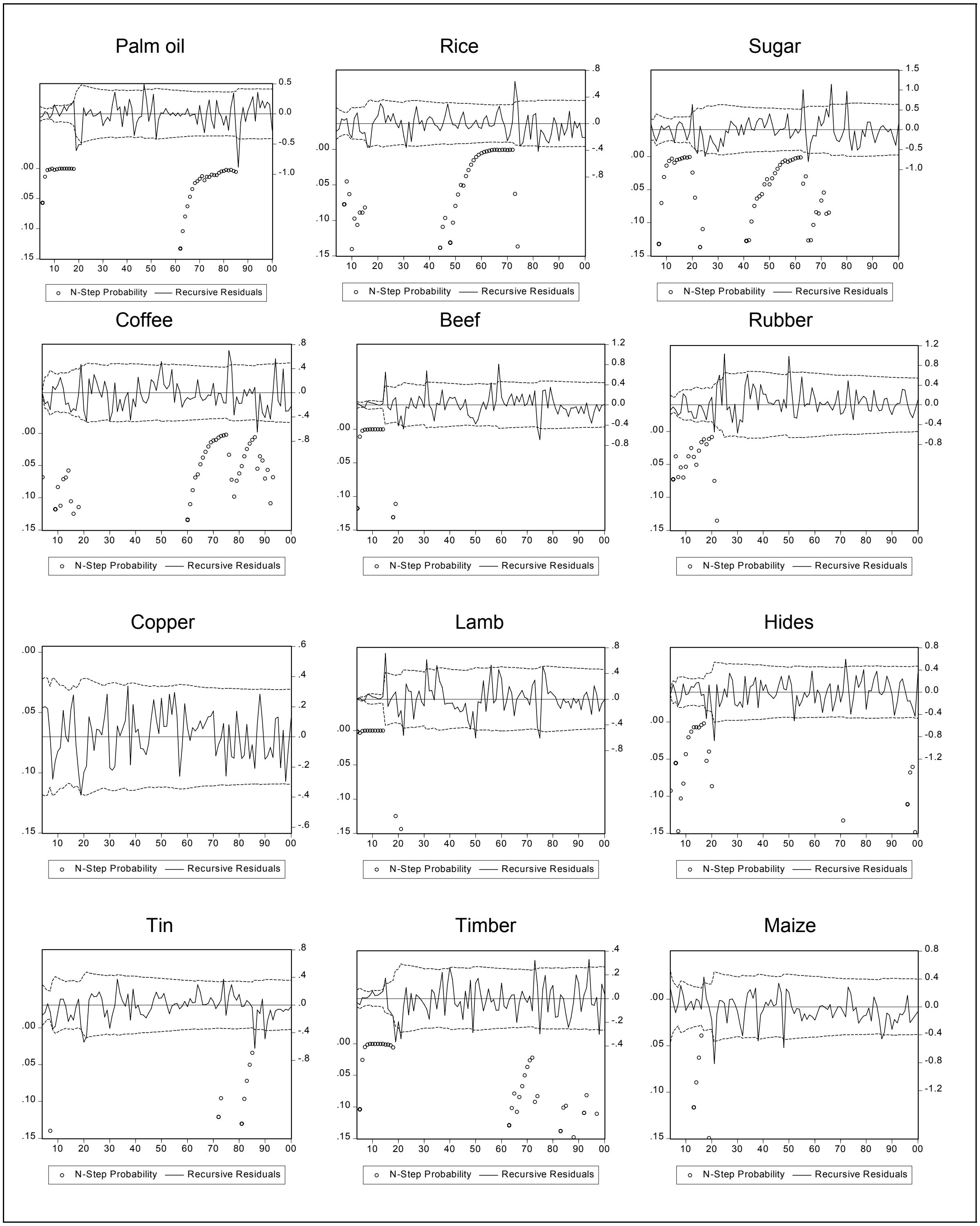
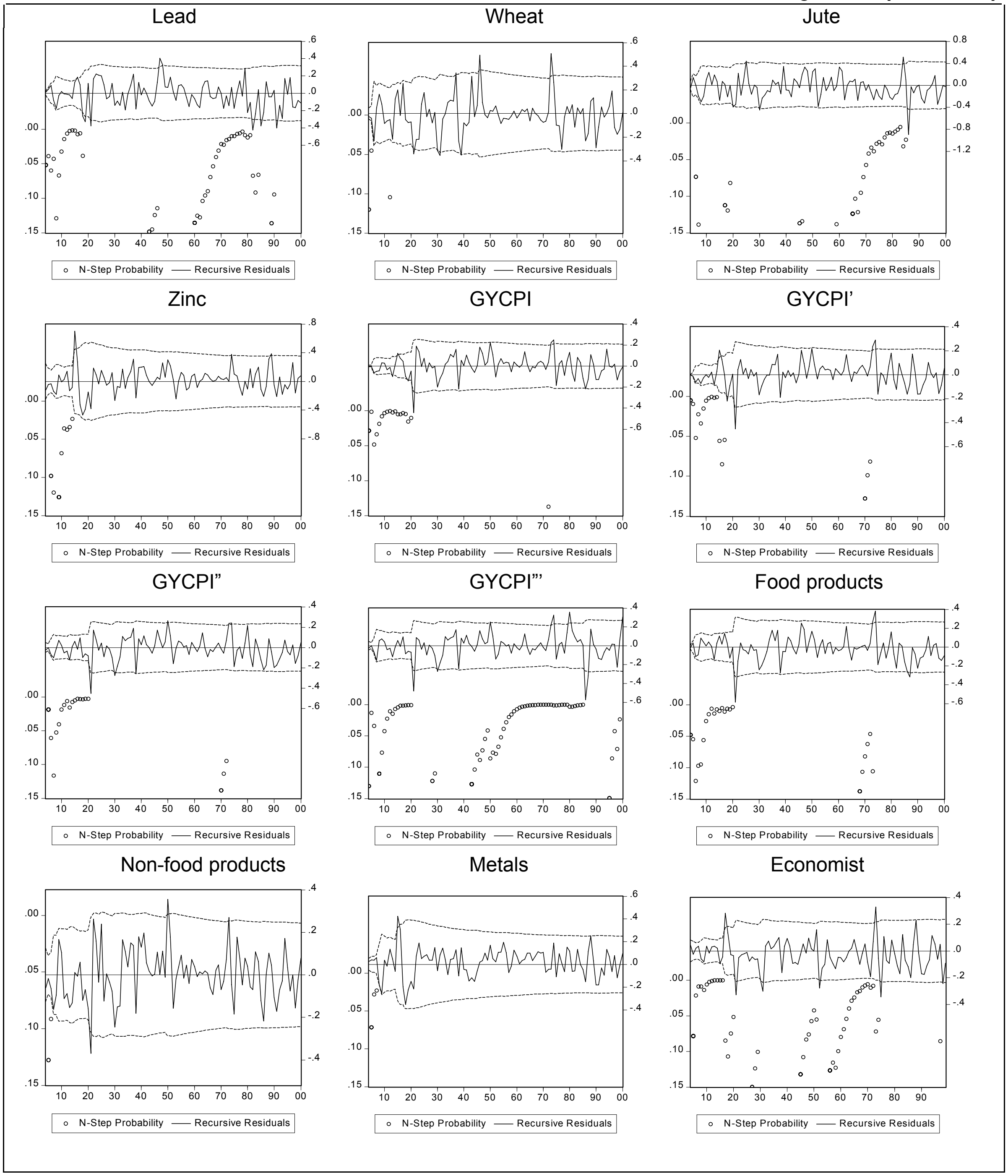


Figure A.2 (concluded)



Source: Calculations of the authors.

Appendix 4

Table A.3

PERRON TEST FOR AN ENDOGENOUS DETERMINATION OF THE DATE OF A STRUCTURAL BREAK

	Innovational outlier (Intercept) IO1 1900-2000 RU	Innovational outlier (Intercept and slope) IO2 1900-2000 RU	Additive outlier (Trend) AO 1900-2000 RU
Aluminium	1938	1940	1948[b]
Bananas	1923	1943	1941[a]
Beef	1957	1957	1982
Cocoa	1945	1965	1998[a]
Coffee	1985	1947	1974
Copper	1951[a]	1951	1925
Cotton	1983	1944	1965
Hides	1950[b]	1916[b]	1905[b]
Jute	1944	1958	1971
Lamb	1945	1945	1926
Lead	1978[a]	1945[b]	1973[b]
Maize	1984	1971	1963
Palm oil	1917[b]	1983[c]	1991[b]
Rice	1988	1971[b]	1975[b]
Rubber	1915[b]	1934[b]	1926[a]
Silver	1960	1971	1933
Sugar	1979[b]	1979	1983
Tea	1983	1951	1962
Timber	1912	1920	1913
Tin	1984	1972	1985
Tobacco	1915	1946	1964
Wheat	1941[c]	1928[b]	1910[c]
Wool	1972	1946[c]	1953[c]
Zinc	1920[c]	1920[a]	1928[c]
GYCPI	1944[b]	1944[a]	1978[a]
GYCPI'	1948[b]	1971[b]	1929
GYCPI"	1983[b]	1970	1978[b]
GYCPI"'	1928	1977	1994
Metals	1915[b]	1951[a]	1940[b]
Food products	1983[b]	1984	1977[a]
Non-food products	1948[b]	1938[b]	1929[a]
ECONOMIST	1915[c]	1915[c]	1920[c]

Source: Calculations of the authors.

[a] 90% significance.
[b] 95% significance.
[c] 99% significance.

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