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Prebisch-Singer Redux

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Motivation

DEVELOPMENT ECONOMISTS HAVE LONG DEBATED whether developing countries should be as specialized as they are in the production and export of primary commodities. Nowhere has this question been debated more hotly than in Latin America. Indeed, it was Latin America that provided the motivation for the seminal contribution of Prebisch (1950) on this topic. He, along with Singer (1950), argued that specialization in primary commodities, combined with a relatively slow rate of technical progress in the primary sector and an adverse trend in the commodity terms of trade,? had caused developing economies to lag behind the industrialized world. Prebisch concluded that, "since prices do not keep pace with productivity, industrialization is the only means by which the Latin American countries may fully obtain the advantages of technical progress." Debate over the validity of Prebisch and Singer's claims, as well as the appropriate policy response, has occupied the literature ever since. really?

While much has happened in Latin America since 1950, the concern about specialization remains as topical as ever. According to noted economic historian and political economist Rosemary Thorp of Oxford University, "The 1990s already saw a return to a primary-exporting role for Latin America. All the signals are that the world economy will push Latin America even more strongly in this direction in the new century, especially in the fields of oil and mining. It behooves us to look very coldly at the political economy and social dimensions of such a model, with more than half an eye on the past. We need to be alert to what will need to change if primary-resource-based growth is to be compatible with long-term economic and social development."

In light of this ongoing concern about commodity specialization in Latin America, we believe it is important to revisit Prebisch's concern of more than 50 years ago that, over the long term, declining terms of trade would frustrate the development goals of the region. This paper has two main objectives. The first is to clarify the issues raised by Prebisch and Singer as they relate to the commodity specialization of developing countries (and Latin America in particular). The second is to reconsider empirically the issue of trends in commodity prices, using recent data and techniques.

The Prebisch-Singer Hypothesis

The Prebisch-Singer hypothesis normally refers to the claim that the relative price of primary commodities in terms of manufactures shows a downward trend.[1] However, as noted earlier, Prebisch and Singer were concerned about the more general issue of a rising per capita income gap between industrialized and developing countries and its relationship to international trade. They argued that international specialization along the lines of "static" comparative advantage had excluded developing countries from the fruits of technical progress that had so enriched the industrialized world.

They rested their case on three stylized facts: first, that developing countries were indeed highly specialized in the production and export of primary commodities; second, that technical progress was concentrated mainly in industry; and third, that the relative price of primary commodities in terms of manufactures had fallen steadily since the late 19th century. Together, these facts suggested that, because of their specialization in primary commodities, developing countries had obtained little benefit from industrial technical progress, either directly, through higher productivity, or indirectly, through improved terms of trade.[2]

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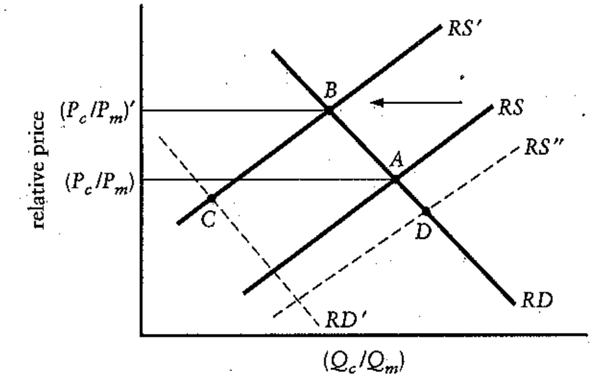
To see this point more clearly, consider figure 5.1, which offers a simple model of the world market for two goods, primary commodities and manufactures. The vertical axis measures the relative price of primary commodities in terms of manufactures, $P_c\ P_m$, while the horizontal axis measures relative quantities, the total quantity of commodities sold on the world market divided by the total quantity of manufactures. The intersection of the relative demand (*RD*) and relative supply (*RS*) schedules determines the world market equilibrium.

If technical progress in the manufacturing sector exceeds that of the primary sector (as Prebisch and Singer supposed), then we should see the supply of manufactures growing faster than the supply of commodities. This would correspond to a declining relative supply of commodities, and this would be represented by a shift to the left of the *RS* schedule to *RS′*. The result would be a shift in the equilibrium from point *A* to point *B* and an increase in the relative price of primary commodities. This relative price change would constitute an improvement in the terms of trade of commodity exporters (whom Prebisch and Singer supposed were developing countries). What we have then is a mechanism, essentially Ricardian in origin, by which technical progress in industrialized countries translates into welfare gains for developing countries.

Figure 5.1 World Market for Primary Commodities Relative to Manufactures



Source: Authors.

The main point of Prebisch and Singer was that this mechanism didn't work: instead of rising, the relative price of commodities in terms of manufactures had actually fallen. They based this conclusion on a visual inspection of the net barter terms of trade—the relative price of exports to imports—of the United Kingdom from 1876 to 1947. The inverse of this was taken to be a proxy for the relative price of primary commodities to manufactures.

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Prebisch and Singer also offered theories as to why the downward trend had occurred and why it was likely to continue. These can be understood by way of figure 5.1 as well. There are essentially two reasons why commodities might experience declining relative prices, despite their lagging technology. One is that something else may prevent the relative supply schedule from shifting to the left or even cause it to shift to the right, like *RS″*. The latter would result in an equilibrium at point *D*, with a lower relative commodity price. The second possibility is that something causes the relative demand schedule to shift to the left (*RD′*) along with relative supply. If the

depletion should ⇒ Left shift

shift from *RD* to *RD′* is greater than that from *RS* to *RS′*, the result would be an equilibrium like point *C*, again with a lower relative commodity price. Over these two alternative explanations for the decline in commodity prices, one involving supply, the other demand, Prebisch and Singer parted company.

Prebisch offered a supply-side theory, based on asymmetries between industrial and developing countries and Keynesian nominal rigidities. The idea was that strong labor unions in industrialized countries caused wages in manufacturing to ratchet upwards with each business cycle, because wages rise during upswings but are sticky during downswings. This, in turn, ratchets up the cost of manufactures. In developing countries, Prebisch argued, weak unions fail to obtain the same wage increases during upswings and cannot prevent wage cuts during downswings. Thus, the cost of primary commodities rises by less than manufactures during upswings and falls by more during downswings, creating a continuous decline in the relative cost of primary commodities, that is, rightward movement in the relative supply schedule.

Singer focused more on the demand side, considering mainly price and income elasticities. Singer argued that monopoly power in manufactures prevented technical progress in that sector from lowering prices, that is, prevented the leftward shift in RS, much like the argument of Prebisch. However, Singer also argued that the demand for primary commodities showed relatively low income elasticity, so income growth tended to lower the relative demand for, and hence relative price of, primary commodities. Moreover, he argued that technical progress in manufacturing tended to be raw-material saving (for example, through the use of synthetics), thereby causing the demand for primary products to grow slower than for manufactures. Both of these arguments would be reflected in a leftward shift in *RD* in figure 5.1.

depends on price

Finally, Prebisch and Singer drew policy implications from what they had found. Both argued that, as the way out of their dilemma, developing countries should foster industrialization. While they stopped short of advocating protectionism, it is clear that they had in mind changing the pattern of comparative advantage. Thus, whether intentionally or not, Prebisch and Singer provided intellectual support for the import substitution policies that prevailed in many developing countries through the 1970s.

The Prebisch-Singer thesis raises a number of questions that we plan to address in this chapter. First, is it reasonable to equate the relative price of commodities with the terms of trade of developing countries in general, and Latin American countries in particular? Second, has the relative price of commodities really declined over the years? Third, are the theories of commodity price determination that Prebisch and Singer put forth plausible? Finally, what policy measures, if any, should developing countries consider toward commodities?

In answering these questions, we shall draw mainly from the literature, although a complete review would be a huge task. For more extensive literature reviews, see Spraos (1980), Diakosavvas and Scandizzo (1991), and Hadass and Williamson (2001). The next two sections discuss the importance of commodity prices for developing countries and some of the factors that determine commodity prices, respectively. This is followed by a brief summary of some new empirical results on the time trend in the commodity terms of trade.[3]

How Important Are Commodity Prices for Developing Countries?

Prebisch and Singer assumed that developing countries were specialized in primary commodities and industrialized countries were specialized in manufactures. This generalization led them to treat the relative price of commodities in terms of manufactures as equivalent to the terms of trade of developing countries (and its inverse, terms of trade of industrialized countries). Of course, developing countries do not export only primary commodities, nor do industrialized countries export only manufactures, and thus commodity prices are distinct from the terms of trade. In this section, we consider the relevance of this distinction.

The fact that industrialized countries do not export only manufactures was addressed early on by Meier and Baldwin (1957) who pointed out that many primary commodities, such as wheat, beef, wool, cotton, and sugar, are heavily exported by industrialized countries. Indeed, Diakosavvas and Scandizzo (1991) note that the developing-country share of agricultural primary commodities was only 30 percent in 1983, down from 40 percent in 1955. Yet Spraos (1980) argues that this fact is immaterial, because the same trends that are observed in the broad index of primary commodity prices are found in a narrower index that includes only developing-country products.

How specialized are developing countries in primary commodities? One way to get at this is to measure the share of commodities in developing-country exports. This is not a perfect measure, however, because it will tend to fluctuate along with relative commodity prices. In particular, if commodity prices are declining, then the value share of commodities in a country's exports may fall, even without any changes in that country's export volume. Bearing in mind this limitation, we look at export shares to get a sense of the degree of specialization and the products in question.

Table 5.1 from Cashin, Liang, and McDermott (2000) shows the commodities that account for a large share of the export earnings for various developing countries. The countries that derive 50 percent or more of their export earnings from a single commodity tend to be in the Middle East and Africa, and the commodity is usually oil. Venezuela is the only such country in Latin America. Several countries receive 20–49 percent of export earnings from a single primary commodity. In Latin America, this includes Chile

Table 5.1 Commodities with a Large Share of Export Earnings in a Given Country (based on annual average export shares, 1992–97)

	50 percent or more of export earnings	*20–49 percent of export earnings*	*10–19 percent of export earnings*
Middle East			
Crude petroleum	Bahrain; Iran, Islamic Rep. of; Iraq; Kuwait; Libya; Oman; Qatar; Saudi Arabia; Yemen, Republic of	Syrian Arab Rep., United Arab Emirates	Egypt, Arab Rep. of
Aluminum			Bahrain
Africa			
Crude petroleum	Angola; Congo, Dem. Rep. of; Gabon; Nigeria	Cameroon, Equatorial Guinea	Algeria
Natural gas		Algeria	
Iron ore		Mauritania	
Copper	Zambia		Congo, Dem. Rep. of
Gold		Ghana, South Africa	Mali, Zimbabwe
Timber (African hardwood)		Equatorial Guinea	Central African Republic, Gabon, Ghana, Swaziland
Cotton		Benin, Chad, Mali, Sudan	Burkina Faso
Tobacco	Malawi	Zimbabwe	
Arabica coffee	Burundi, Ethiopia	Rwanda	
Robusta coffee	Uganda		Cameroon
Cocoa	São Tempe and Principe	Côte d'Ivoire, Ghana	Cameroon
Tea			Kenya, Rwanda
Sugar		Mauritius	Swaziland
Western Hemisphere			
Crude petroleum	Venezuela, R. B. de	Ecuador, Trinidad and Tobago	Colombia, Mexico
Copper		Chile	Peru
Western Hemisphere (continued)			
Gold			Guyana
Cotton			Paraguay
Arabica coffee			Colombia, El Salvador, Guatemala, Honduras, Nicaragua
Sugar		Guyana, St. Kitts and Nevis	Belize
Bananas		Honduras, St. Vincent	Costa Rica, Ecuador, St. Lucia
Fishmeal			Peru
Rice			Guyana
Europe, Asia, and Pacific			
Crude petroleum		Azerbaijan, Brunei Darussalam, Norway, Papua New Guinea, Russian Federation	Indonesia, Kazakhstan, Vietnam
Natural gas	Turkmenistan		
Aluminum		Tajikistan	
Copper		Mongolia	Kazakhstan, Papua New Guinea
Gold		Papua New Guinea	Uzbekistan
Timber (Asian hardwood)		Lao PDR, Solomon Islands	Cambodia, Indonesia, Myanmar, Papua New Guinea
Timber (softwood)			Latvia, New Zealand
Copra and coconut oil	Kiribati		
Cotton		Pakistan, Uzbekistan	Azerbaijan, Tajikistan, Turkmenistan

Source: Cashin, Liang, and McDermott 2000.

in copper, and several others in bananas and sugar. Still more countries have primary-export-revenue shares in the 10–19 percent range.

Table 5.2 shows the top two exported primary commodities (along with the export shares of these commodities) for several Latin American countries over the past century. Since 1900, the export share of the top two primary commodities has fallen in every country except República Bolivariana de Venezuela. Even in República Bolivariana de Venezuela, it has fallen since 1950. Today only three countries, Chile, Cuba, and República Bolivariana de Venezuela have commodity export shares above 40 percent. This decline may be simply because of declining commodity prices, but more likely it reflects changing comparative advantage: developing countries are competitive in certain areas of manufacturing, while industrialized countries have moved into the production of services. It may also reflect the effect of import-substitution policies in developing countries over the latter half of the century.

Several studies have taken a more rigorous approach to measuring the importance of commodity prices for the terms of trade of developing countries. Bleaney and Greenaway (1993), for example, estimate a cointegrating regression for nonoil developing countries from 1955–89, in which terms of trade of the developing countries (from IMF data) is expressed as a log-linear function of an index of commodity prices and real oil prices. The results show that the series are cointegrated, and that for every 1 percent decline in the relative price of commodities there is a 0.3 percent decline in the terms of trade of nonoil developing countries. These results are similar to those of Grilli and Yang (1988) and Powell (1991).

By far the most comprehensive study on this topic is Bidarkota and Crucini (2000). They take a disaggregated approach, examining the relationship between the terms of trade of 65 countries and the relative prices of their major commodity exports. Bidarkota and Crucini find that at least 50 percent of the annual variation in national terms of trade of a typical developing country can be accounted for by variation in the international prices of three or fewer primary commodity exports.

In the final analysis, the importance of commodities in developing countries depends on the precise question one wishes to address. Commodity price trends and fluctuations are clearly important to any policy designed to stabilize commodity prices or the income of commodity producers, such as a stabilization fund or commodity agreement. As noted by Cuddington and Urzúa (1989) and Deaton and Laroque (1992), the effectiveness of a stabilization fund depends crucially on whether shocks to commodity prices are temporary or permanent. Further, an understanding of commodity price trends should also inform longer-term policies affecting the allocation of productive factors across sectors, as was the original intent of Prebisch and Singer.[4] In both of these instances, however, the more disaggregated the data, the better.

Beyond informing policy, Prebisch and Singer sought to use their theory to explain the performance gap between developing and industrialized

Table 5.2 Top Two Commodities Exported by Latin American Countries, 1900–95 (share of each commodity (%) in total exports)

	1900	*1910*	*1920*	*1930*	*1940*	*1950*	*1960*	*1970*	*1980*	*1990*	*1995*
Argentina	wool (24) wheat (19)	wheat (23) wool (15)	wheat (24) meat (18)	wheat (19) meat (18)	meat (23) wheat (16)	wheat (17) meat (15)	meat (22) wool (14)	meat (25) wheat (6)	meat (13) wheat (10)	meat (7) wheat (6)	oil (8) wheat (5)
Bolivia	silver (39) tin (27)	tin (54) rubber (16)	tin (68) silver (11)	tin (84) copper (4)	tin (80) silver (6)	tin (67) lead (9)	tin (66) lead (7)	tin (50) gas (16)	tin (43) gas (25)	gas (26) zinc (16)	zinc (11) gas (10)
Brazil	coffee (57) rubber (20)	coffee (51) rubber (31)	coffee (55) cocoa (4)	coffee (68) cotton (3)	coffee (34) cotton (18)	coffee (62) cocoa (7)	coffee (55) cocoa (6)	coffee (32) iron (7)	soya (12) coffee (10)	soya (9) iron (8)	soya (8) iron (6)
Chile	nitrate (65) copper (14)	nitrate (67) copper (7)	nitrate (54) copper (12)	nitrate (43) copper (37)	copper (57) nitrate (19)	copper (52) nitrate (22)	copper (67) nitrate (7)	copper (79) iron (6)	copper (46) iron (4)	copper (46) fish (4)	copper (39) wood (6)
Colombia	coffee (49) gold (17)	coffee (39) gold (16)	coffee (62) gold (13)	coffee (64) oil (13)	coffee (62) oil (13)	coffee (72) oil (13)	coffee (75) oil (13)	coffee (59) oil (13)	coffee (54) oil (13)	oil (23) coffee (21)	coffee (20) oil (19)
Costa Rica	coffee (60) banana (31)	banana (53) coffee (32)	coffee (51) banana (33)	coffee (67) banana (25)	coffee (54) banana (28)	coffee (56) banana (30)	coffee (53) banana (24)	coffee (29) banana (29)	coffee (27) banana (22)	banana (24) coffee (17)	banana (24) coffee (14)

(continued)

Table 5.2 Top Two Commodities Exported by Latin American Countries, 1900–95 *(continued)*
(share of each commodity (%) in total exports)

	1900	*1910*	*1920*	*1930*	*1940*	*1950*	*1960*	*1970*	*1980*	*1990*	*1995*
Cuba	sugar (61) tobacco (23)	sugar (70) tobacco (24)	sugar (87) tobacco (10)	sugar (68) tobacco (17)	sugar (70) tobacco (8)	sugar (82) tobacco (5)	sugar (73) tobacco (8)	sugar (75) tobacco (4)	sugar (82) nickel (5)	sugar (74) nickel (7)	sugar (50) nickel (22)
Mexico	silver (44) copper (8)	silver (28) gold (16)	oil (67) silver (17)	silver (15) oil (14)	silver (14) zinc (13)	cotton (17) lead (12)	cotton (23) coffee (9)	cotton (8) coffee (5)	oil (65) coffee (4)	oil (32) coffee (2)	oil (10) — —
Peru	sugar (25) silver (18)	copper (20) sugar (19)	sugar (35) cotton (26)	oil (33) copper (21)	oil (26) cotton (21)	cotton (34) sugar (15)	cotton (18) copper (17)	fish (27) copper (25)	oil (20) copper (18)	copper (18) fish (13)	copper (19) fish (15)
Uruguay	wool (29) hides (28)	wool (40) hides (23)	wool (40) meat (30)	meat (37) wool (27)	wool (45) meat (22)	wool (48) meat (19)	wool (57) meat (20)	wool (32) meat (32)	wool (17) meat (17)	wool (16) meat (11)	meat (14) wool (9)
Venezuela, R. B. de	coffee (43) cacao (20)	coffee (53) cacao (18)	coffee (42) cacao (18)	oil (82) coffee (10)	oil (88) coffee (3)	oil (94) coffee (1)	oil (88) iron (6)	oil (87) iron (6)	oil (90) iron (2)	oil (79) aluminum (4)	oil (75) aluminum (4)

Source: Thorp 1998, 347.
Note: — = not available.

countries. For this purpose, it is more important to understand the terms of trade of developing countries than to understand commodity prices. This is the approach taken by Hadass and Williamson (2001). They bypass the question of the relationship between the terms of trade and commodity prices altogether and simply reexamine evidence on the Prebisch-Singer hypothesis, using country-specific terms of trade data, instead of commodity-price data. They construct estimates of the terms of trade for 19 countries, developing and industrialized, and aggregate these into four regions: land-scarce Europe, land-scarce Third World, land-abundant New World, and land-abundant Third World. Simply by comparing averages, they find that the terms of trade improved for all regions except the land-scarce Third World. They argue that this is due in part to rapidly declining transport costs during the sample period, which is consistent with Ellsworth's (1956) criticism of Prebisch and Singer.

trans costs
↓ ⇒ ∇ cost
adv for
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Determinants of Commodity Prices: What Explains the Relative Price of Primary Commodities?

While most of the literature on the Prebisch-Singer hypothesis has focused on testing the claim of declining relative commodity prices, several papers attempt direct tests of the theories put forth by Prebisch and Singer. Diakosavvas and Scandizzo (1991) examine Prebisch's theory of asymmetrical nominal rigidities. In particular, they examine the implication that during upswings, the prices of primary products and manufactures should move roughly in tandem, while in downswings, prices of primary products should fall much more than do those of manufactures. They test this by looking at whether the elasticity of primary product prices with respect to manufactures prices is higher on downswings than on upswings. It turns out that the data reject the hypothesis for all but five commodities (nonfood, rice, cotton, rubber, and copper).

Bloch and Sapsford (1997, 2000) estimate a structural model to assess the contribution to commodity prices of a number of factors described by Prebisch and Singer. They build a model that assumes marginal cost pricing in the primary sector and markup pricing in manufactures. Wages are explicitly introduced to try and pick up the effects of unions in manufactures. The model also allows for biased technical change in á la Singer.

Recognizing the potential nonstationarity of the series, Bloch and Sapsford first difference the entire model and apply a two-stage least squares procedure. (While this has the intended effect of producing stationarity, it also has the unfortunate effect of sweeping out long-run relationships between the variables.) Bloch and Sapsford (1997) find that the main contributing factor to declining commodity prices is raw-material-saving technical change. There is also some contribution from faster wage growth in manufactures and a steadily increasing manufacturing markup.

The manufacturing markup interpretation is suspect, however, as the markup is based on price minus labor and intermediate input costs, leaving out rents to other factors, such as capital and land.

Whereas Bloch and Sapsford focus on microeconomic factors affecting commodity prices, Borensztein and Reinhart (1994) and Hua (1998) focus on macroeconomic determinants. Borensztein and Reinhart (1994) construct a simple model where commodities are used as inputs in the production of manufactures, and their prices are quoted in U.S. dollars on world markets. Global commodity demand, therefore, depends positively on world production of manufactures and negatively on the U.S. dollar real exchange rate. As the dollar appreciates in real terms, the relative price of commodities in non-U.S. industrial countries rises, thereby choking off their demand for commodity inputs. The authors assume market clearing where commodity demand is equated to an exogenous commodity supply, and they proceed to estimate both supply and demand effects.

As in Bloch and Sapsford, the model is first differenced before estimation by generalized least squares (GLS). When estimated without the supply component, the model fits well until the mid-1980s, after which it vastly overpredicts the relative price of commodities. The fit is restored, however, once supply shocks are introduced, and it is improved still further after account is taken of the fall in industrial production in Eastern Europe and the former Soviet Union in the late 1980s.

Hua (1998) estimates a demand-side model of commodity prices, similar to that of Borensztein and Reinhart, but he adds in the real interest rate (to capture the opportunity cost of holding commodities) and lagged oil prices. He estimates the model using a reduced-form error-correction specification. He finds that the hypothesis of a stationary long-run relationship between commodity prices and the levels of industrial output and real exchange rate cannot be rejected.

Empirical Evidence on Trends in Primary Commodity Prices: Is There a Downward Trend in the Relative Price of Commodities?

Evidence Up through Grilli-Yang (1988)

The bulk of the empirical literature on the Prebisch-Singer hypothesis looks for a secular decline in the relative price of primary commodities in terms of manufactures, rather than directly at the terms of trade of developing countries. Until fairly recently, the largest single obstacle to this search was a lack of good data. Prebisch and Singer had based their conclusions on the net barter terms of the United Kingdom from 1876 to 1947. Subsequent authors criticized the use of these data on several grounds, and various attempts were made to correct for data inadequacies.

Spraos (1980) discusses these criticisms in detail (see box 5.1 for a summary) and also provides estimates based on data that are marginally better than those used by other authors up to that point. Spraos concluded that over the period 1871–1938 a deteriorating trend was still detectable in the data, but its magnitude was smaller than suggested by Prebisch and Singer. When the data was extended to 1970, however, the trend became statistically insignificant. Implicit in this conclusion is the notion that the parameters of the simple time-trend model have not remained constant over time. We return to this point later.

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Sapsford (1985) extended the Spraos data and considered the possibility of a once-and-for-all (or "structural") break in the time trend of relative commodity prices. He showed there to be a significant overall downward trend of 1.3 percent per year with a large, upward, nearly parallel, shift in the trend line around 1950. ?

Many of the data issues raised by early authors were put to rest by Grilli and Yang (1988), who carefully constructed a price index of 24 internationally traded nonfuel commodities spanning the period 1900–86. The nominal prices are drawn from a World Bank database consisting of annual observations on the 24 nonfuel commodities, as well as two energy commodities: oil and coal. The latter are not included in the Grilli and Yang (hereafter referred to as "GY") index. The nonfuel group includes 11 food commodities: bananas, beef, cocoa, coffee, lamb, maize, palm oil,

Box 5.1: Bad Data?

Numerous authors criticized Prebisch and Singer's use of British terms of trade data to proxy for relative commodity prices. Here are the four main problems, according to Spraos (1980) and references therein:

1) Britain's terms of trade were not representative of the terms of trade of industrialized countries on the whole.
2) Industrialized countries export primary commodities, too, so the inverse of their terms of trade is a bad measure of relative commodity prices.
3) British exports were valued without transport costs, while its imports were valued inclusive of transport costs. Thus, declining transport costs alone could improve the British terms of trade, thereby overstating the drop in commodity prices.
4) Introducing new manufactured goods and improving the quality of existing ones may push up the price index of manufactures, giving the impression of a decline in the relative price of commodities.

rice, sugar, tea, and wheat; seven nonfood agricultural commodities: cotton, hides, jute, rubber, timber, tobacco, and wool; and six metals: aluminum, copper, lead, silver, tin, and zinc. Based on 1977–79 shares, these products account for about 54 percent of the world's nonfuel commodity trade (49 percent of all food products, 83 percent of all nonfood agricultural products, and 45 percent of all metals).

To construct their nominal commodity-price index, Grilli and Yang weighted the 24 nominal prices by their respective shares in 1977–79 world commodity trade. To get a real index, GY divided their nominal commodity-price index by a manufacturing-unit-value (MUV) index, which reflects the unit values of manufactured goods exported from industrial countries to developing countries.[5] This is a natural choice of deflators, given Prebisch and Singer's (hereafter referred to as "PS") concern about the possibility of a secular deterioration in the relative price of primary commodity exports from developing countries in terms of manufacturing goods from the industrial world.

The MUV-deflated GY series, which has recently been extended through 1998 by IMF staff economists, is shown in figure 5.2.[6]

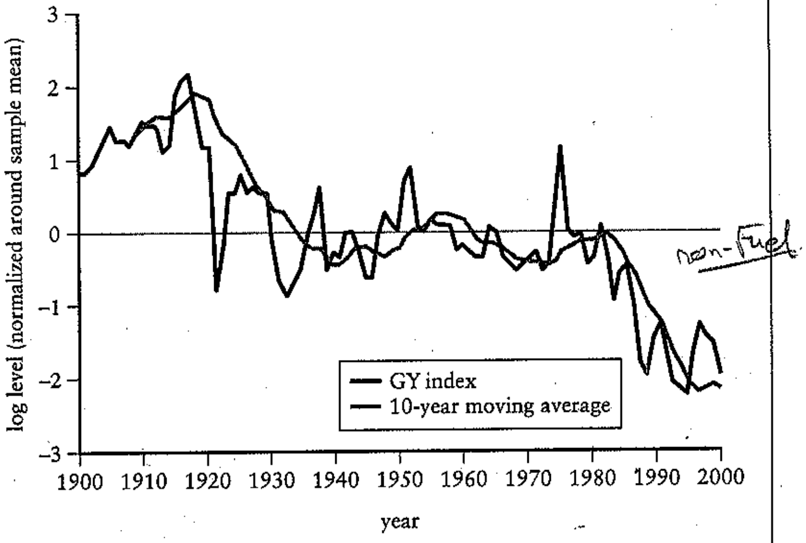
Using their newly constructed index, which covered the 1900–86 period, Grilli and Yang estimated a log-linear time trend and found a significant downward trend of –0.6 percent per year, after allowing for the presence of a downward break in the level of the series in 1921. They, therefore, concluded that their findings supported the PS hypothesis.

Post Grilli-Yang Work: Econometric Issues

Since the publication of the GY paper and associated long-span dataset in late 1980s, there has been a resurgence in empirical work on long-term trends in commodity prices. The search for a secular trend has shifted from the issue of data quality to econometric issues involved in estimated growth rates or trends in nonstationary time series. Most authors have used the GY dataset, extended to include more recent data in many cases. In a recent paper, Cashin and McDermott (2002) use The Economist's index of industrial commodity prices covering an even longer time span: 1862–1999 or 140 years! They find a downward trend of –1.3 percent per year.

Visual inspection of the MUV-deflated GY series in figure 5.2, as well as its 10-year moving average, leaves one with the strong impression that it has trended downward over time, as PS conjectured. Modern time-series econometrics, however, has taught us that it is potentially misleading to assess long-term trends by inspecting time plots or estimating simple log-linear time trend models (see box 5.2). Although the GY series in figure 5.2 does *not* appear to be mean stationary, it is critical to determine the source of nonstationarity before attempting to make inferences about the presence of any trend. Possible sources of nonstationarity are:

Figure 5.2 Grilli-Yang Commodity Price Index Deflated by the MUV



Source: Authors' calculations.

- A deterministic time trend
- A unit root process, with or without drift[7]
- One or more structural breaks in the mean or trend of the univariate process
- General parameter instability in the underlying univariate model

The key econometric issues are, in short, the possible presence of *unit roots* and *parameter instability* in the univariate models being estimated. To facilitate a discussion of these issues and to put the existing literature into context, we first specify a general log-linear time trend model that may or may not have a unit root. Second, we describe three types of structural breaks in this framework, where there are sudden shifts in model

Box 5.2: Unit Root Perils

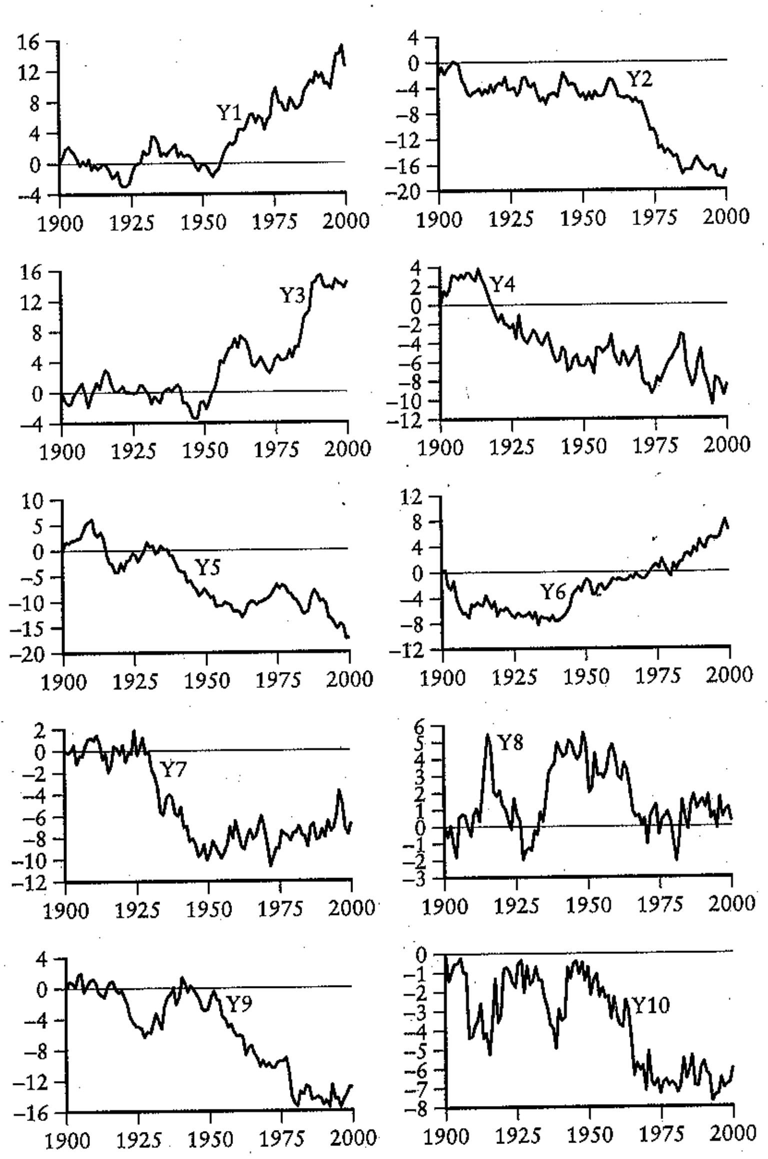
It is now well known in the time series econometrics literature that attempting to assess long-run trends and detect structural breaks based on graphical evidence and time series models is a highly misleading exercise, especially if the time series are, in fact, unit root process. To illustrate, consider the 10 series shown in the following box figures. Which series exhibit clear positive or negative trends? Which series show structural breaks? Which series have pronounced cyclical behavior?

Reviewing your answers to these three questions, you may find it somewhat surprising to learn that each of the 10 series is a driftless random walk. So, despite appearances, none of these series has any deterministic trend, cyclical component, or structural break(s)!

Even though these series are really driftless random walks, if you regress each of the series on a constant and a time trend (and correct for apparent first-order serial correlation in the residuals), you will (incorrectly) conclude that 9 of the 10 series have statistically significant time trends—6 are significantly negative; 3 are significantly positive. This is an example of the *spurious regression* phenomenon highlighted by Granger and Newbold (1974). There is also spurious cyclicity, reflected in the form of spuriously "significant" serial correlation coefficients (see Nelson and Kang (1981)). Finally, if you eyeball the data to identify dates when there have apparently been structural breaks, then add dummy variables (at the point where visual inspection suggests that the series "breaks") to your log-linear trend models, you will undoubtedly find spuriously significant structural breaks as well.

It is true that visual inspection of the deflated GY series in the box figures leaves little doubt that it is nonstationary in the mean, but this need not be the result of a deterministic time trend like (5.1). The random walk process above is the simplest example of a time series that is nonstationary in the mean due to the presence of a unit root. Unit root processes, with or without drift, are also nonstationary. The time series and unit root possibilities are nested neatly within the specification in (5.1)–(5.2). If $\rho < 0$, and $\beta \neq 0$, we have a deterministic time trend model. PS predicts $\beta < 0$. If $\rho = 1$ and $\beta \neq 0$, we have a unit root process with drift. Again, if $\beta < 0$, this is consistent with the PS hypothesis. If $\rho = 1$ and $\beta = 0$, we have a driftless unit root process. If real commodity prices are characterized by a unit root, this might be of concern to developing countries or to others who specialize or are contemplating greater specialization in primary commodities, but not for the reasons PS articulated. The concern would have to be refocused on managing risk, rather than on coping with secular deterioration.

Box 5.2: Unit Root Perils *(continued)*



parameters. A more general type of parameter instability, where parameters are hypothesized to follow random walks, is briefly summarized.

Trend Stationary vs. Difference Stationary Models: Unit Roots

Attempts to estimate the long-term growth rate or trend in an economic time series typically begin with a log-linear time trend model:

$$\ln(y_t) = \alpha + \beta \cdot t + \varepsilon_t \tag{5.1}$$

In the PS literature, $y = P_C P_M$ is the ratio of the aggregate commodity price index to the manufacturing goods unit value. The coefficient β of the time index t is the (exponential) growth rate; it indicates the rate of improvement ($\beta > 0$) or deterioration ($\beta < 0$) in the relative commodity price y_t. It is important to allow for possible serial correlation in the error term ε_t in (5.1). Econometrically, this improves the efficiency of the parameter estimation; economically, it captures the often-pronounced cyclical fluctuations of commodity prices around their long-run trend.

The error process in (5.1) is assumed to be a general autoregressive, moving average (ARMA) process:

$\varepsilon_t = \sum \rho \varepsilon_{t-i} + \sum u_{t-i}$

$$(1-\rho L)A(L)\varepsilon_t = B(L)u_t \tag{5.2}$$

It will be convenient in what follows to factor the autoregressive component of the error process in a way that isolates the largest root in the AR part of the error process; this root is denoted ρ. The terms $(1 - \rho L)A(L)$ and $B(L)$ are AR and MA lag polynomials, respectively. The innovations u_t in (5.2) are assumed to be white noise.

long-term

A critical issue will be whether $|\rho| < 1$, temporary indicating that the error process is stationary, or whether $\rho = 1$, permanent indicating nonstationarity due to the presence of a unit root over time. In the former case, (1)-(2) is referred to as the *trend stationary (TS) model,* indicating that fluctuations of y_t around its deterministic trend line are stationary. y_t itself, however, is nonstationary unless $\beta = 0$.

If, however, y_t (or equivalently the error process in (5.2)) contains a unit root, estimating the TS model—with or without allowance for (supposed) structural breaks—will produce spurious estimates of the trend (as well as spurious cycles). An appropriate strategy for estimating the trend β in this case is to first-difference the model (5.1)–(5.2) to achieve stationarity. The result is the so-called *difference stationary (DS) model,* a specification in terms of growth rates rather than log-levels of the y_t series:

TS= control for t = ok
DS: 1st diff = ok

$$(1-L)\ln(y_t) \equiv D\ln(y_t) = \beta + v_t \tag{5.3}$$

where L and D are the lag and difference operators, respectively. The error term in (5.3) follows an ARMA process:

$$A(L)v_t = B(L)u_t \tag{5.4}$$

In the DS model, a significant negative estimate of the constant term, β, would be support for the PS hypothesis.

Using the extended GY dataset (1900–98), suppose we ignore the possibilities of unit roots and structural breaks and simply estimate the TS model. The following results are obtained:

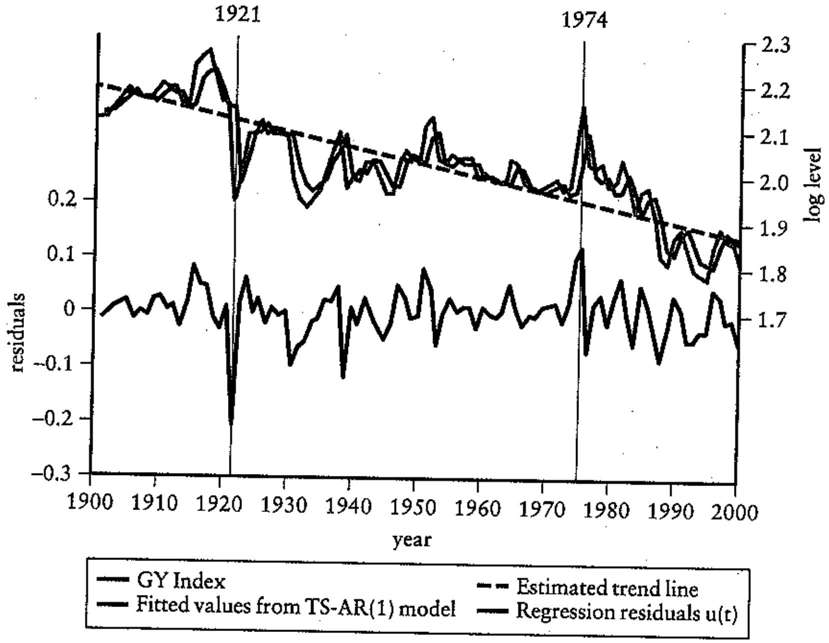
$$y_t = 2.19 - 0.003t + \varepsilon_t$$
$$\text{where } \varepsilon_t = 0.74\,\varepsilon_{t-1} + u_t$$

The error process is adequately modeled as a first-order AR process. There is a statistically significant trend coefficient equal to –0.3 percent per year ($t = -5.23$). Fitted values from the TS model, the long-term trend estimate, and the regression residuals are shown in figure 5.3. The figure reveals some potential problems. First, the fitted regression line does not fit the data especially well. Note that the fitted line consistently lags the turning points in the actual data.

Moreover, the residuals have possible outliers at 1921 and to a lesser extent in 1974 (or 1973). Reexamining the GY series itself in light of these observations, one might speculate that there have been structural breaks in 1921 and 1974. More formal methods for identifying the timing of a possible break (or two) are considered next. These methods indicate clear evidence of a break in 1921, with a second, but statistically insignificant, break in the early 1970s or mid-1980s.

One way to assess the structural stability of the TS-AR(1) model is to calculate recursive residuals and the two-standard error bands for the hypothesis that the recursive residuals come from the same distribution as those from the estimated model. As seen in figure 5.4, the recursive residuals in 1921 and 1974 are "large," suggesting structural breaks. Figure 5.4 also shows p-values for an N-step forecast test for each possible forecast sample. To calculate the p-value for 1920, for example, one would use data from 1900 through 1920 to estimate a TS-AR(1) model. This model is then used to forecast $y(t)$ for the remaining N years of the sample: 1921–98. A test statistic that incorporates the forecast errors, comparing the forecast with the actual value, for the N-steps ahead can be constructed to test the null hypothesis that such forecast errors could have been obtained from the underlying TS-AR(1) model with no structural

Figure 5.3 A Secular Deterioration in Real Commodity Prices?

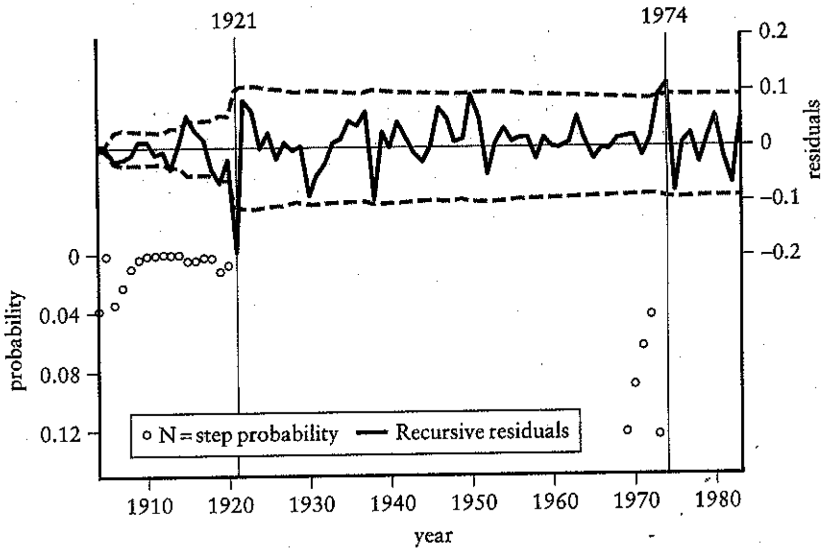


Source: Authors' calculations.

break. The p-value for the null hypothesis of no structural break gives the probability of finding an even larger test statistic if the null is, in fact, true. If the p-value is smaller than the size of the test, typically 0.01 or 0.05, then one should reject the null hypothesis of no structural breaks.

As seen in figure 5.4, the p-values very near 0.4 in the 1910–20 period indicate that the test statistic is so large that the probability of finding a larger one under the null is virtually zero. That is, this graph clearly shows that if the model is fitted with pre-1921 data and used to forecast into the future, there is clear rejection of parameter stability. If, instead, one uses data up through the 1940s or 1950s or 1960s, on the other hand, parameter stability is not rejected. If one uses data through the early 1970s to forecast commodity prices through the end of the 1990s, there is again instability—albeit somewhat less severe (judging from the p-values on the left-hand scale of the graph).

Figure 5.4 Evidence of Parameter Instability in TS-AR(1) Model



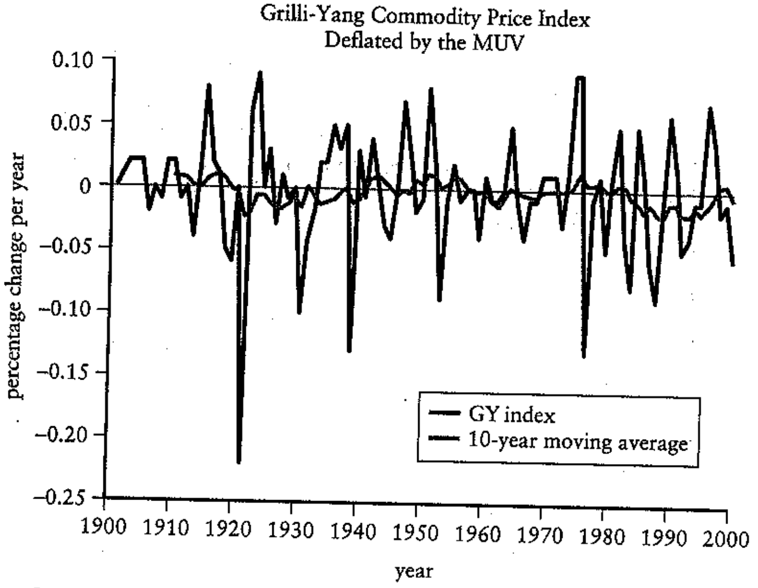
Source: Authors' calculations.

This evidence indicates that the issue of structural breaks or parameter instability must be taken seriously if one chooses the TS model for analyzing the long-term trends in primary commodity prices.

Consider now the DS model, which uses first-differences of the logged real commodity price series shown in figure 5.5, to estimate the growth rate in commodity prices. This specification is appropriate if one believes that the GY series is a unit root process.

Note that the $D(y)$ series is very volatile. The 10-year moving average is, not surprisingly, much smoother. It also "goes through the data" much better than it did the 10-year moving average of the log-levels in figure 5.2. This is consistent with the presumption that $D(y)$ is stationary, but y is not. The average value of $D(y)$ is small and negative, −0.3 percent per year (including the huge −22.0 percent outlier in 1921). Given the high variance of the series, however, it is not surprising that the null hypothesis of a zero growth rate cannot be rejected.

Figure 5.5 A Volatile Unit Root Process?



Source: Authors' calculations.

The regression results presented in table 5.3 are for a DS model, with two lags of DGY being sufficient to eliminate serial correlation in the residuals.

The recursive residual and N-step ahead forecast analysis, shown in figure 5.6, again suggests that there is a structural break in 1921. With the DS model, however, 1921 appears to be the only troublesome episode.

What is clear up to this point? In sum, the possibility of finding statistical significance for the trend in the real GY commodity price index depends critically on whether one believes *a priori*, or concludes, on the basis of unit root tests, that Grilli-Yang is trend stationary, or whether it contains a unit root. Regardless of whether the TS or DS specification is chosen, there is evidence that one or two breaks or parameter instability may be a problem.

Structural Breaks and Parameter Instability

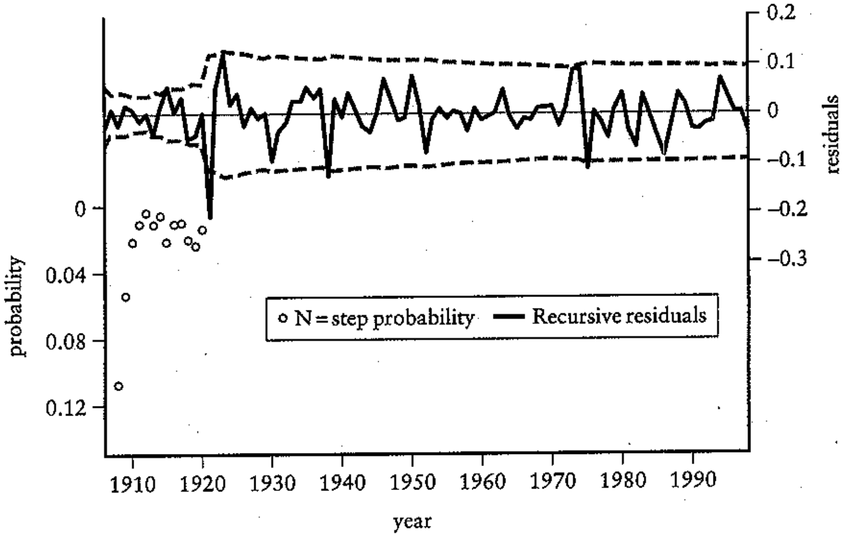
It has long been recognized that estimated parameters in models like the TS and DS models earlier will be biased, or even meaningless, if the true

Table 5.3 Estimation Results for a Difference Stationary Model for the GY Series

Dependent variable: DGY *Sample (adjusted): 1903–98*	
Constant	−0.004
	(0.005)
DGY(-1)	0.004
	(0.101)
DGY(-2)	−0.259
	(0.101)
R^2	0.066
Observations	96

Source: Authors' calculations.
Note: Standard errors are given within parentheses.

Figure 5.6 Evidence of Parameter Instability in DS Model



Source: Authors' calculations.

parameters do not remain constant over time. Suppose, for example, that the true growth rate was –4.0 percent in the first half of the sample, but 2.0 percent in the second half. An econometrician who ignored the shift in parameters might incorrectly conclude that the growth rate was a uniform –2.0 percent over the entire sample.

Trend Δ vs cycles?

To consider the possibility of a change in parameters (α, β) in the TS model or β in the DS model,[8] one typically constructs a dummy variable: $DUM_{TB} = 0$ for all $t < TB$ and $DUM_{TB} = 1$ for all $t \geq TB$ where TB is the hypothesized break date. Using this "level-shift" dummy, as well as its first difference (a "spike" dummy) and a dummy-time trend interaction term, yields the "TS with break" model and the "DS with break" model, respectively:

TS with break model

$$\ln(y_t) = \alpha_1 + \alpha_2 DUM_{TB} + \beta_1 t + \beta_2 (t - TB) \cdot DUM_{TB} + \varepsilon_t \qquad (5.5)$$

DS with break model

$$D(\ln(y_t)) = \alpha_2 D(DUM_{TB}) + \beta_1 + \beta_2 \cdot DUM_{TB} + \nu_t \qquad (5.6)$$

These specifications are general enough to encompass the three types of breaks described in Perron's (1989) classic paper on testing for unit roots in the presence of structural breaks (which will be discussed later). His model A ("crash" model) involves only an abrupt shift in the level of the series; that is, $\alpha_2 \neq 0$, $\beta_2 = 0$. In model B ("breaking trend" model), there is a change in the growth rate, but no abrupt level shift: $\alpha_2 = 0$, $\beta_2 \neq 0$. Finally, model C ("combined" model) has change in both the level and growth rate: $\alpha_2 \neq 0$, $\beta_2 \neq 0$.

Suppose that one knows *a priori*, or decides on the basis of unit root testing, whether the TS or DS specification is appropriate. Then, if the break date, TB, is assumed to be known, it is straightforward to test for the presence of structural breaks by examining the T-statistics on α_2 and β_2. A test for a break of type C could be carried out using a $\chi^2(2)$ test for the joint hypothesis that $\alpha_2 = 0$ and $\beta_2 = 0$.

The latter is equivalent to (one variant of) the well-known Chow test for a structural break. More recent work on tests for parameter stability warns against arguing that the break date TB is known. Andrews (1993); Ploberger, Kramer, and Kontrus (1989); and Hansen (1992, 2001), for example, develop methods for testing for the presence of a possible structural break at an unknown date using algorithms that search over all possible break dates.

Recently, there have been attempts in the macroeconomics literature to extend the unknown break date literature to consider *two* break points at unknown dates (see, for example, Lumsdaine and Papell (1997) and Mehl (2000)). An obvious issue that this extension raises is: why only two breaks rather than, say, three or four?

Authors developing parameter stability tests have also considered the alternative hypothesis where the parameters are assumed to follow a random walk. In this case, the model parameters are generally unstable, in a way that cannot be captured by a one-time shift at any particular date. This test of general parameter stability is a good diagnostic test when assessing the adequacy of a particular model specification.

Hansen (1992, 321) provides an excellent overview of the issue and possible approaches to dealing with it:

> One potential problem with time series regression models is that the estimated parameters may change over time. A form of model misspecification, parameter nonconstancy, may have severe consequences on inference if left undetected. In consequence, many applied econometricians routinely apply tests for parameter change. The most common test is the sample split or Chow test (Chow 1960). This test is simple to apply, and the distribution theory is well developed. The test is crippled, however, by the need to specify *a priori* the timing of the (one-time) structural change that occurs under the alternative. It is hard to see how any non-arbitrary choice can be made independently of the data. In practice, the selection of the breakpoint is chosen either with historical events in mind or after time series plots have been examined. This implies that the breakpoint is selected conditional on the data and therefore conventional critical values are invalid. One can only conclude that inferences may be misleading.

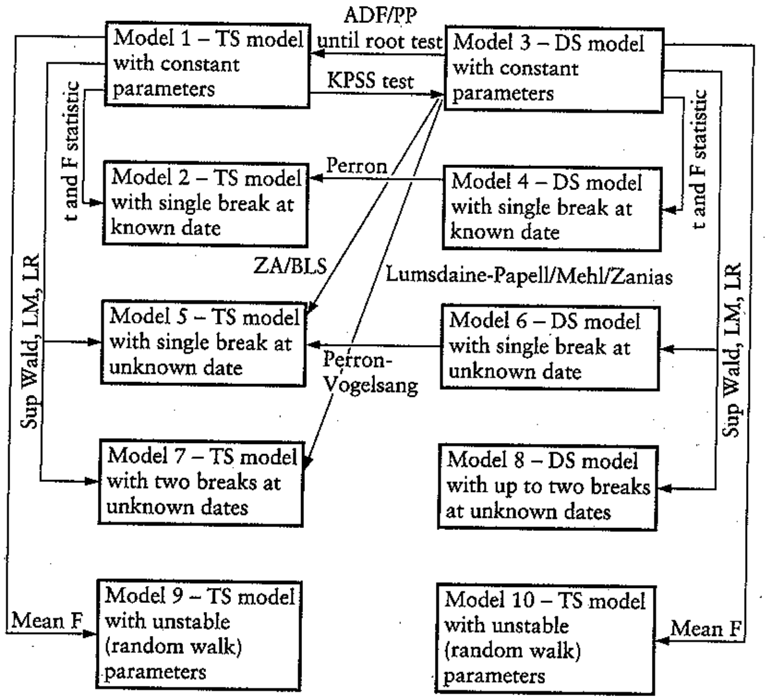
An alternative testing procedure was proposed by Quandt (1960), who suggested specifying the alternative hypothesis as a single structural break of unknown timing. Until recently, a difficulty with Quandt's test was that the distributional theory for the test statistic was unknown. This problem, however, was independently solved by Andrews (1993), Chu (1989), and Hansen (1990).

In situations where one is tempted to argue that there are several structural breaks, it probably makes sense to ask whether the situation might be better described as one of general parameter instability.

A Selective Review of Post Grilli-Yang Empirical Work

As mentioned earlier, the literature through Grilli-Yang (1988) used the TS model—as indicated by model 1 in figure 5.7, which summarizes approaches taken in the literature—to estimate the long-term trend in real commodity prices. A number of these authors recognized the possibility of structural changes in the form of one-time shifts in the level or trend in the

Figure 5.7 Alternative Specifications



Source: Authors' summary of the literature.

real commodity price series. That is, they compared models 1 and 2. For example, Sapsford (1985) found a break in 1950 using pre-Grilli-Yang data, as mentioned earlier. Grilli and Yang (1988) and Cuddington and Urzúa (hereafter referred to as "CU") (1989) both identified a breakpoint in 1921 using the Grilli-Yang dataset for the period 1900–83. Contrary to Grilli-Yang, CU argued that, after accounting for the highly significant downward shift in the level of the real Grilli-Yang price index in 1921, the trends on either side of the break were not significantly different from zero in the TS specification. Not surprisingly, if one ignored the one-time downward step in the data, the estimated trend coefficient β appears to be negative and significant. This illustrates the potential for incorrect statistical inferences if structural shifts are ignored.

CU also demonstrated that the structural break in 1950 detected by Sapsford (1985) using the trend stationary model on pre-Grilli-Yang data was not significant when using the Grilli-Yang data once the 1921 break was included.

CU (1989) were the first to carry out unit root tests on the Grilli-Yang commodity price index. They were unable to reject the unit root hypothesis, and they, therefore, stated a preference for DS models rather than TS models when estimating the long-term trend in real commodity prices. Using data from 1900–83, they were unable to reject the null hypothesis that $\beta = 0$ in the DS model in (3)-(4), where β is the long-term drift in real commodity prices. This finding was robust to the inclusion or exclusion of the one-time drop in the level of the Grilli-Yang series in 1921. In terms of figure 5.7, CU (1989) considered models 3 and 4, and they formally tested model 3 against model 1 (assuming no unit root) and model 3 against model 4 (assuming there is a unit root). By carrying out augmented Dickey-Fuller (ADF) tests (1979), they compared model 3 to model 1, and using a new unit root test in Perron (1989), they tested model 4 against model 2.

Applying ADF unit root tests as well as Perron-ADF tests that allow for a possible structural break at a predetermined break date, CU (1989) showed that the unit root hypothesis cannot be rejected for the Grilli-Yang index. When CU estimated the DS model using Grilli-Yang data from 1900–83, the estimated long-term growth rate was statistically insignificant, regardless of whether one included a spike dummy to account for the downward shift in the level of the real Grilli-Yang series in 1921.

The DS specification, therefore, leads to the conclusion that real commodity prices follow a driftless unit root process. The policy implications from this specification are quite different from those based on the CU's TS model with a one-time level shift in 1921. The risk entailed for commodity producers, exporters, and commodity stabilization fund managers is considerably greater if one believes that the true model is the DS specification. The CU unit root tests failed to reject the null hypothesis of a unit root, but such tests have notoriously low power, so no definitive conclusion is warranted.

Note that the DS model with a one-time level shift in 1921 is a very plausible candidate model for the Grilli-Yang series. In fact, it is the specification preferred by Cuddington and Urzúa (1989). The year 1921, moreover, occurs early in the sample, precisely the situation where Leybourne, Mills, and Newbold (1998) warn that DF tests are likely to lead to false rejections of the unit root hypothesis when the true data-generating process is unit root with a structural break! In spite of this bias, CU did not reject the unit root when they assumed a known break date. Assuming an unknown break date implies smaller (that is, more negative) critical values for the resulting Zivot-Andrews-Perron (ZAP)-ADF test. So, again, one would not expect to reject the unit root hypothesis.

Cuddington (1992) repeated the exercise of testing for unit roots (with or without breaks at possible break dates determined by visual inspection) for each of the 24 component commodities in the Grilli-Yang index (1900–83). Some commodities had unit roots; others did not. Some commodities had negative price trends, while others had positive trends. Surprisingly, not a single commodity had a structural break in 1921![9] This led Cuddington and Wei (hereafter referred to as "CW") (1992) to conjecture that there was some aggregation issue involved in the construction of the Grilli-Yang index, as theirs was an arithmetic index. Cuddington and Wei constructed a geometric index, so that the results from the individual commodities should be reflected in the geometric index, as it was just a simple weighted average of the logs of the individual commodity prices that comprise the index. Using the CW index (over the slightly extended period 1900–88), they found that unit root tests are inconclusive. The estimated trend in the real commodity price index, however, turned out to be statistically insignificant regardless of whether one used the TS or DS model specification.

Subsequent work has reconsidered Cuddington and Urzúa's claim of a trendless series with a break in 1921. Powell (1991), for example, found three downward jumps, in 1921, 1938, and 1975, and no continuous trend. Ardeni and Wright (1992) used a "trend plus cycle model" and extend the Grilli-Yang data to 1988 to find a continuous trend between −0.14 percent and −1.06 percent, depending on the exact model specification. Moreover, this trend survives with or without a structural break in 1921. Bleaney and Greenaway (1993) avoided the issue of a structural break in 1921 by considering 1925–91 data, and they instead found a downward jump in 1980, with no continuous trend.

León and Soto (1997) and Zanias (2005) applied the Zivot-Andrews/Banerjee, Lumsdaine, and Stock (ZA/BLS) method for testing for unit roots in the presence of a single break at an unknown break point. Zanias, in particular, found that this method identifies 1984 as the primary break point. It is, however, difficult to know how to interpret a break point in a portion of the sample that Andrews and others recommend should be trimmed off, because it is too close to the end of the sample. Zanias went on to reapply the ZA/BLS approach to find a second break, conditional on the presence of the first break in 1984. This sequential procedure chose 1921 as the second break point

Although the PS literature has extensively explored the possibility of structural breaks, the more general phenomenon of parameter instability has only recently been explored. See Cuddington, Ludema, and Jayasuriya (2002) (hereafter referred to as "CLJ"), who apply Hansen's approach to the Grilli-Yang commodity index. Apart from the econometric issues raised by, for example, Hansen's quote earlier, parameter instability has interesting implications for testing the PS hypothesis. PS did not claim that the long-run trend would necessarily remain constant over time, only that it would be negative!

A New Look at Growth Rates, Possible Breaks, and Unit Root Tests

In testing the PS hypothesis, our primary interest is in the growth rate β in the deflated Grilli-Yang index. Has it been negative as PS predicted? Has it been relatively stable over time? Or has this parameter shifted or drifted over time, or exhibited a sharp structural break or breaks? In our particular application, we are less interested in the presence or absence of unit roots per se than was the applied macroeconometric literature. Unfortunately, it is difficult to estimate the growth rate β without making a decision on the presence or absence of a unit root first. Ideally, we would also like to formally test for the presence of structural breaks without prejudging the case of whether the series has a unit root. This objective, however, appears to be beyond our reach at this time.

The strategy in CLJ (2002) is the following. First estimate augmented ZAP-ADF-like regressions allowing for, at most, two structural breaks at unknown dates. Having searched for the two most plausible break dates, test whether each break is statistically significant. If both breaks are significant, assume two breaks in what follows. If only one break is statistically significant, reestimate the ZAP-ADF equation with a single break at an unknown date and test to see whether the remaining break is statistically significant. The results of this exhaustive unit root in the presence of two possible breaks at unknown dates is, alas, rather inconclusive—not in identifying the likely break points, but in resolving the issue of whether there is a unit root in the real commodity price index.

Given the uncertainty surrounding the question of unit roots, CLJ estimate both TS and DS models with one or two breaks. The diagnostic tests suggest that if the TS model is adopted, two break points are detected—in 1921 and 1985. With the DS specification, however, only a single break—in 1921—appears statistically significant. These specifications are chosen by carrying out a grid search over all possible pairs of break dates.

Estimated TS and DS Models with Two Breaks

Next, we consider the TS and DS models in turn, using our search algorithm to choose the dating of two break points.[10] As discussed earlier, we need to include only the level-shift and time-interaction dummies to allow for breaks of type A, B, and C in the TS model. Thus the criterion for choosing the break dates (TB1, TB2) is the sup$\chi^2(4)$ statistic from the set of all $\chi^2(4)$ statistics testing the joint significance of the two dummies associated with all possible pairs of break dates. Analogously, in the DS specification, we need to include only the spike and level-shift dummies. The criterion is again a sup$\chi^2(4)$ statistic.

Once the two most plausible break points have been identified in the TS and DS specifications, respectively, there are three subsamples of the

Grilli-Yang index to consider. It is necessary to estimate the growth rates for each segment: pre-TB1, TB1 through TB2, and post-TB2. Estimates of the trend segments for both the TS and DS specifications are shown in table 5.4. Also reported is the Wald test of the null hypothesis that each trend coefficient is equal to zero. A rejection of the hypothesis indicates the presence of a significant trend in the respective subperiod.

Examining the table, we find that sup$\chi^2(4)$ statistics for both the TS and DS specifications are "large" (relative to the standard 1 percent critical value for $\chi^2(4)$ of 13.28). The mean$\chi^2(4)$ statistic for the DS model is very small, suggesting no issue of general parameter instability. The mean$\chi^2(4)$ statistic for the TS model is close enough to the standard critical value that it is impossible to guess the outcome of a formal parameter stability test based on simulated critical values.

The TS model estimation places the two breaks in 1921 and 1985. Moreover, the $\chi^2(2)$_TB1 and $\chi^2(2)$_TB2 stats for 1921 and 1985, respectively, are similar in magnitude, with 1985 being slightly larger (14.36 vs. 13.25, whereas the 1 percent critical value for $\chi^2(2) = 9.21$).

The resulting calculations for the TS model growth rates and their χ^2 statistics (conventional p values noted) indicate that the trend in all three subperiods are not statistically different from zero. In conclusion, therefore, if one rejects the unit root hypothesis and accepts the TS model, the Grilli-Yang series is best characterized as a zero-growth series that has experienced two significant downward level shifts (type A breaks), first in 1921 and then again in 1985.

Table 5.5 shows the estimation results for the best-fitting TS model with two breaks. Figure 5.8 shows the actual logged GY series, the fitted values and residuals from the best-fitting TS specification with two breaks, and the forecasted values starting in 1900 in order to show the long-run trend segments more clearly. The tests summarized in table 5.4, indicate that the trend is insignificantly different from zero in each of the three segments of the TS model: pre-1920, 1921–84, and post-1984.

In contrast to the TS model, the DS model identifies the two break years as 1921 and 1974, rather than 1985. Note that for the DS model, the sup$\chi^2(4)$ is very large, while the mean$\chi^2(4)$ statistic is quite small. (For comparison, the standard $\chi^2(4) = 13.28$.) Also, the 1921 break has a much higher $\chi^2(2)$ stat than the 1974 break. Together, these χ^2 statistics suggest that, if one uses the DS specification, the Grilli-Yang series is well characterized by one (1921) or possibly two (1921 and 1974) structural breaks rather than general parameter instability. Examining the $\chi^2(2)$_TB1 (= 19.32) and the $\chi^2(2)$_TB2 (= 4.77) statistics, it is clear that the 1921 break is significant, while the 1974 break is not statistically significant.[11] Thus, the DS specification requires only a single break in 1921.[12]

Table 5.4 Grid Search Results for Two Possible Breaks at Unknown Dates (TB1, TB2)[a]

Type of model	*TS model*	*DS model*
Type of structural break dummies	*level and time interaction*	*level and spike*
Chosen break points TB1 and TB2	**1921 and 1985**	**1921 and 1974**
Sup$\chi^2(4)$	34.43	47.40
Mean$\chi^2(4)$	8.07	3.35
(Segmented) trend[b]		
1. pre_TB1	0.003 (0.184)	0.003 (0.656)
2. TB1 through TB2	−0.001 (0.130)	0.000 (0.970)
3. post_TB2	−0.002 (0.587)	−0.011 (0.031)
χ^2 stat(2)_TB1	13.25	19.32
χ^2 stat(2)_TB2	14.36	4.77

Source: Authors' calculations.

Note: a. On a Pentium III processor, the grid search program to consider all break date pairs runs for approximately 20 minutes each for the TS and DS models. In both cases, the maximum number of lags of the dependent variable considered (k) was six.

b. The p-value for the hypothesis that the trend coefficient is equal to zero is given in parentheses. P-values that are higher than your chosen test size (say .05) indicate failure to reject the null hypothesis of a zero trend for the given segment of the data. These p-values ignore the fact that TB1 and TB2 were chosen so as to maximize sup$\chi^2(4)$. Thus the p-values on the trend segments are possibly inaccurate.

Estimated DS Models with a Single Break

We now search for a single break in the Grilli-Yang series using the DS model. In this case, we include only the level and spike dummies in the estimation. We now use the sup$\chi^2(2)$ statistic to test the hypothesis that these two dummies are zero. Figure 5.9 graphs the $\chi^2(2)$ for the DS model.

Here, the maximum sup$\chi^2(2)$ has a value of 32.26 and occurs in 1921. The second highest sup$\chi^2(2)$ has a value of 6.28 and occurs in 1975. In addition, the mean$\chi^2(2)$ statistic is 1.58, a contrastingly low value compared to either the supχ^2 or the 1 percent critical value of 9.21 from the standard $\chi^2(2)$ distribution. Therefore, with the DS specification, a single downward level shift in 1921 but with no ongoing (stochastic) trend fits the data well.

Table 5.5 Estimation Results for a Trend Stationary Model with Two Breaks for the Grilli-Yang Series

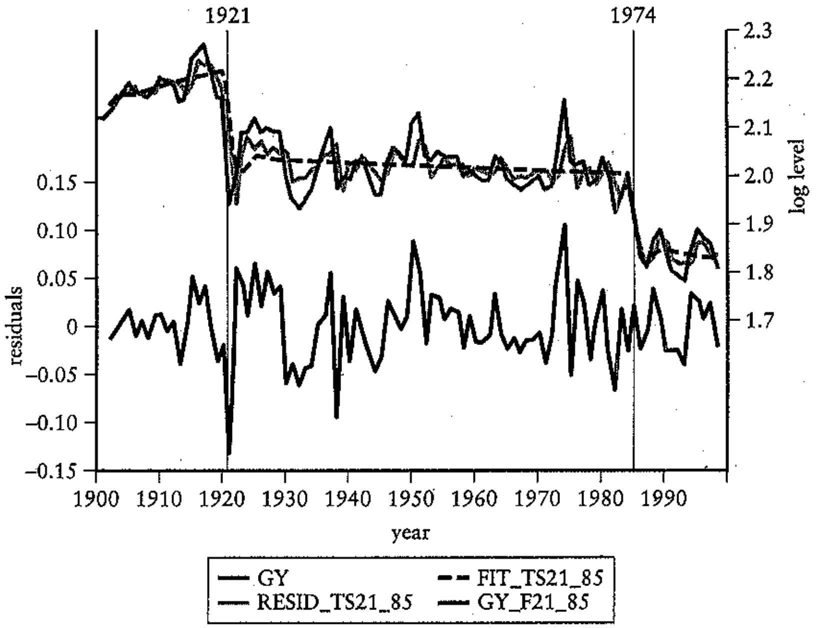
Dependent variable: GY *Sample (adjusted): 1902–98*	
Constant	1.489
	(0.203)
GY(−1)	0.622
	(0.098)
GY(−2)	−0.314
	(0.096)
TREND	0.002
	(0.002)
DUM1921	−0.069
	(0.026)
DUM1921*TREND	−0.003
	(0.002)
DUM1985	−0.012
	(0.244)
DUM1985*TREND	−0.001
	(0.003)
R^2	0.881
Observations	97

Source: Authors' calculations.
Note: Standard errors are in parentheses.

Conclusions

Despite 50 years of empirical testing of the Prebisch-Singer hypothesis, a long-run downward trend in real commodity prices remains elusive. Previous studies have generated a range of conclusions, due in part to differences in data but mainly due to differences in specification, as to the stationarity of the error process and the number, timing, and nature of structural breaks. In this chapter, we have attempted to allow the data to tell us the proper specification. In our most general specification (model 8, in figure 5.7), which allows for a unit root and searches for two structural breaks of any kind, we find the most likely pair of breaks to be in 1921 and 1974, but the 1974 break is statistically insignificant. Moreover, we cannot reject the hypothesis of a unit root. If we search for only one structural break, we find one very clearly in 1921, again, with no rejection of the unit root hypothesis. This model indicates also that there is no drift, either positive or negative, before or after 1921.

Figure 5.8 A Segmented Trend Stationary Model?



Source: Authors' calculations.

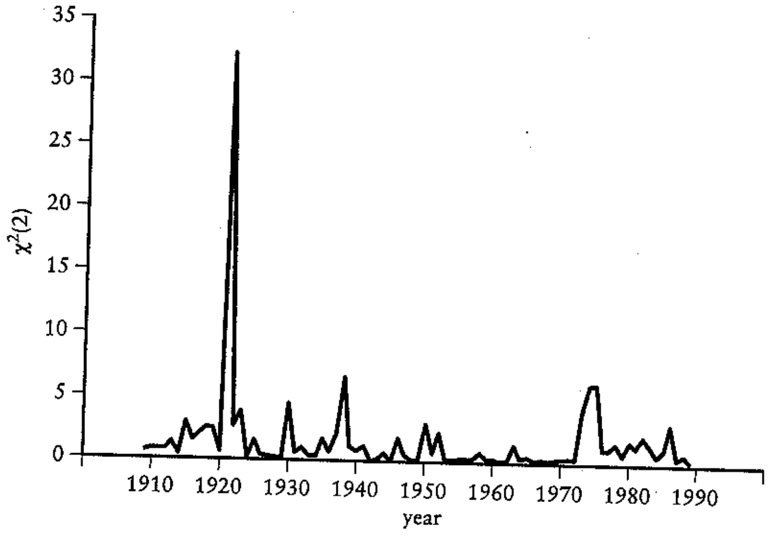
downward break in 1921 no trend before or after

If we assume the Grilli-Yang series is trend stationary, we find much fuzzier results. The two-break model (model 7, in figure 5.7) puts the breaks in 1921 and 1985, with both breaks borderline significant. The three segments in this case (before, between, and after the breaks), show no trend. The model with one break puts the break in 1946, but is rejected in favor of model 1 (TS with no break). Only in the case of model 1—the model studied by researchers since the beginning of Prebisch-Singer testing—can one find a significant negative trend. Yet model 1 is inconsistent with our results in N-step ahead forecasting.

We conclude that the preponderance of evidence suggests that the series is well characterized as a unit root process with a single level-shift break (type A) in 1921.

temporary shocks

Figure 5.9 The $\chi^2(2)$ Sequence for DS Model with One Break



Source: Authors' calculations.

Notes

*The authors would like to thank Shuichiro Nishioka and two anonymous reviewers for helpful comments.

1. There is some debate as to whether the Prebisch-Singer hypothesis refers narrowly to a prediction about relative commodity prices or more generally to the idea that commodity specialization is inimical to development. We take the narrower view, mainly because Prebisch and Singer's argument hinges so crucially on their prediction about commodity prices. Furthermore, there are other arguments against commodity specialization in the literature that are unrelated to the terms of trade. It would be misleading to group all of them under the banner of the Prebisch-Singer hypothesis. That said, we strive in this chapter to give Prebisch and Singer their due, for their hypothesis, their reasoning, and what they perceived as its broader implications.

2. Singer (1950) went further to argue that foreign direct investment had also failed to spread the benefits of technical progress, because it tended to be isolated into enclaves with developing countries and, thus, have few spillovers.

3. These results, and the underlying methodology, are discussed in more detail in Cuddington, Ludema, and Jayasuriya (2002).

4. It is not at all clear that a policy of this kind is called for. The point is that, if a policy is to be considered, it should be done based on the information about commodity price trends.

5. GY also considered a U.S. manufacturing price index as a deflator and concluded that their results were not much affected by the choice of deflator.

6. We thank Paul Cashin of the International Monetary Fund Research Department for providing these data.

7. In principle, a series could contain both a deterministic trend and a unit root or more than one unit root; we ignore these cases here.

8. It is also possible to allow for shifts in the model parameters that describe the error process, as well as its serial correlation and variance, but we do not consider this extension here.

9. Cuddington found breaks for only coffee (1950) and oil (1974); the latter is not in the GY index.

10. Each specification requires the inclusion of two dummies for each break date. It can be shown that the break dates must be separated by at least one period to avoid perfect multicollinearity.

11. What about the calculated growth rates for each segment in the DS specification if we assume there are *two* breaks? Results for the DS model are slightly different from those obtained from the TS model. In spite of a statistically insignificant trend in each of the first two subperiods, the DS model identifies the existence of a "possibly significant" *negative* trend of 1.09 percent in the post-1974 period.

12. This is consistent with the ZAP-ADF tests reported in CLJ (2002), which found a single break and were unable to reject the null hypothesis of a unit root.

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