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Keynes, Cocoa, and Copper: In Search of Commodity Currencies

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Keynes, Cocoa, and Copper: In Search of Commodity Currencies[1]

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Abstract

This paper examines whether the real exchange rates of commodity-exporting countries and the real prices of their commodity exports move together over time. Using IMF data on the world prices of 44 commodities and national commodity export shares, we construct new monthly indices of national commodity export prices for 58 countries over 1980-2002. A long-run relationship between real exchange rates and real commodity prices is found for about two-fifths of the commodity-exporting countries. Also, the behavior of the real exchange rate of commodity currencies is found to be independent of the nominal exchange rate regime. The average half-life of adjustment of real exchange rates to commodity-price-augmented purchasing power parity is found to be about eight months, which is much shorter than Rogoff's (1996) consensus estimate of three to five years, and provides an important missing piece of the PPP puzzle.

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I. INTRODUCTION

"The neglect to allow for the effect of changes in the terms of trade is, perhaps, the most unsatisfactory characteristic of Prof. Cassel's "Purchasing Power Parity Theory of the Foreign Exchanges". For this not only upsets the validity of his conclusions over the long period, but renders them even more deceptive over the short period..." (Keynes (1930, p.336)).

Attempts by economists to model long-run movements in real (price-level adjusted) exchange rates have typically proven to be rather unsuccessful. Meese and Rogoff (1983) demonstrated that a variety of linear structural exchange rate models failed to forecast more accurately than a naïve random walk model for both real and nominal exchange rates, and their key finding has not been overturned in the succeeding three decades. If the real exchange rate follows a random walk, then innovations to the real exchange rate persist and the time series can fluctuate without bound. This result is contrary to the theory of purchasing power parity (PPP), which states that there is a constant equilibrium level to which exchange rates converge, such that foreign currencies should possess the same purchasing power. Accordingly, PPP has proven to be a weak model of the long-run real exchange rate, and recent work has emphasized the time-varying nature of the long-run real exchange rate.

There is a large empirical literature on the determinants of the long-run real exchange rate, which has emphasized sectoral productivity differentials, government spending, cumulated current account imbalances, and interest rate differentials as important drivers of long-run deviations from purchasing power parity (see Froot and Rogoff (1995) and Rogoff (1996) for recent surveys). This literature has mainly concentrated on understanding the sources of real exchange rate fluctuations in developed countries, and the fruits of this research have been mixed, with many studies failing to find a statistical link between real exchange rates and the above explanators.[2]

In contrast to the preponderance of developed country studies of the behavior of real exchange rates, evidence on the behavior of developing country real exchange rates has been scarce. Those studies which have examined the determinants of developing country real exchange rates have largely focused on Latin America, and have emphasized the role of movements in the terms of trade in driving real exchange rate movements (see Diaz-Alejandro (1982), Edwards (1989), and Edwards and Savastano (1999)). There is also an extensive literature for some developed countries which links exogenous movements in the terms of trade of commodity-exporting countries and changes in their real exchange rates, particularly for commodity exporters Canada and Australia (see, among others, Amano and van Norden (1995) and Gruen and Wilkinson (1994)).

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[2] See Edison and Melick (1999) on the typical failure to find cointegration between real exchange rates and real interest rate differentials, and Rogoff (1996) on the mixed empirical track record of the Balassa-Samuelson effect on real exchange rates.

Rogoff (1996) summarizes the multitude of potential explanators offered by researchers in their attempts to resolve the PPP puzzle, which concerns the finding of many researchers that the speed of mean reversion of real exchange rates is too slow to be consistent with PPP. Chief among these explanators has been the recognition that real factors have a role in the determination of real exchange rates, through such channels as: the Balassa-Samuelson effect; real interest rate differentials; and portfolio balance models (where higher net foreign assets drive an appreciation of the exchange rate). However, while such factors may have resonance for developed countries, in the context of commodity-exporting countries (almost all of which are also developing countries), it is difficult to see much of a role for these real factors given the slow pace of relative productivity improvements in the production of tradables, the typical presence of both capital controls and underdeveloped domestic financial markets, and limited capital mobility.[3]

For developing countries, because primary commodities dominate their exports, fluctuations in world commodity prices have the potential to explain a large share of movements in their terms of trade. However, while terms of trade fluctuations have been considered a key determinant of developing country real exchange rates (Edwards (1989); De Gregorio and Wolf (1994)), it is surprising that there has been no comprehensive empirical work done to assess the mechanisms through which changes in relative commodity prices affect the real exchange rate.[4] The main purpose of this paper is to systematically examine the relationship between the real exchange rate and relative commodity prices for all commodity-dependent economies.

Importantly, Obstfeld and Rogoff (2000) and Chen and Rogoff (2002) point out that in the presence of sticky producer prices and perfect pass-throughs, standard measures of the terms of trade will move one-to-one mechanically with the real exchange rate, making it extremely difficult to identify causality between the real exchange rate and terms of trade. More generally, if the extent of exchange rate pass-through is less for exports than for imports, a depreciation of the local currency will raise the local currency price of exports relatively less than it will raise the local currency price of imports—this will yield a decline in the terms of trade. Deaton and Miller (1996) used a measure of the terms of trade

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[3] Many low-income countries have underdeveloped financial markets and domestic credit markets with publicly determined interest rates, typically at nonmarket levels (Agénor and Montiel (1996)). This makes the use of nominal (or real) interest differentials as a determinant of real exchange rate movements in developing countries problematic, as domestic interest rates are largely endogenous to domestic monetary policy.

[4] Two earlier country-specific analyses have been the work of Edwards (1985), who examined the relationship between real coffee prices and Colombia's real exchange rate; and that of Chen and Rogoff (2002), who find that commodity prices drive the real exchange rates of developed country commodity exporters Australia, New Zealand, and Canada.

expressed in *world prices* to circumvent this potential endogeneity problem. We follow Deaton and Miller and construct, for each commodity-dependent economy, indices of real commodity prices which are defined as the world (nominal) price of their commodity exports relative to the world price of manufactured goods exports. Our measure of the world price of commodity exports aggregates changes in world commodity prices using actual national export shares of the commodity exports. For large commodity-exporting countries, world relative commodity prices are likely to be better at capturing the exogenous component of terms of trade shocks than standard terms of trade measures (Chen and Rogoff (2002)).[5]

This paper takes Keynes (1930) seriously, and will examine whether relative price movements within the tradables sector, in particular changes in the relative price of commodity exports to imports, are a major determinant of movements in developing country real exchange rates. As shocks to relative commodity export prices are both volatile and persistent, controlling for this important source of real shocks may help resolve some of the major empirical puzzles which persist in modeling real exchange rates. In particular, we will determine how many commodity-exporting countries have 'commodity currencies,' in that movements in their real exchange rates are affected by movements in the real price of their commodity exports.[6]

The key objective of this paper is to examine whether movements in real commodity export prices can explain fluctuations in the real exchange rates of individual commodity-dependent countries. The paper does so in several ways. First, a new monthly dataset of country-specific export price indices is constructed for 58 countries over the period January 1980 to March 2002. Each country's export price index is a geometric weighted average of world commodity prices, using country-specific export shares as weights. Second, using empirical techniques which allow for structural shifts in the long-run relationship between time series, we find strong evidence of a long-run relationship between the real exchange rate and real commodity prices for about two-fifths of the commodity-exporting countries in our sample. For these commodity currencies, movements in real commodity prices are an important determinant of long-run deviations of real exchange rates from purchasing power

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[5] Deaton and Laroque (1992) found that as the terms of trade is an aggregate price index, it is a poor measure of the short-lived booms and long-lived troughs frequently observed in the prices of major exports of commodity-dependent countries. Kouparitsas (1997), Bidarkota and Crucini (2000), and Baxter and Kouparitsas (2000) all find that, for developing countries, real commodity prices (the relative prices of nonfuel commodities to manufactured goods) are much more volatile than the terms of trade.

[6] The term 'commodity currency' has an alternative meaning in the international finance literature. It was earlier associated with post-Second World War proposals for the creation of a commodity-reserve currency, whereby a specified basket of commodities would be designated as the unit of account, with a price fixed in terms of the nominal currency (see Graham (1937)).

parity. Third, qualifying Mussa's (1986) finding for industrial countries, the behavior of the real exchange rate of commodity currencies is found to be independent of the nominal exchange rate regime. Fourth, contrary to the findings of Calvo, Reinhart, and Végh (1995) for Latin America, commodity currencies are not real exchange rate targeters, as movements in the real exchange rate are found to be more volatile than movements in real commodity prices. Fifth, weak exogeneity tests carried out within a vector error correction framework indicate highly significant causality running from real commodity prices to the real exchange rate. When deviations from the long-run equilibrium relationship occur in commodity currencies, it is primarily the real exchange rate that adjusts to restore long-run equilibrium. For commodity currencies, the average half-life of adjustment of real exchange rates to commodity-price-augmented purchasing power parity is about eight months, which is much shorter-lived than Rogoff's (1996) consensus estimate of the half-life of deviations from purchasing power parity of between three to five years. This rapid speed of mean reversion supports the equilibrium approach to exchange rate determination, and provides an important missing piece of the PPP puzzle.

The paper is organized as follows. Section II briefly sets out the theoretical relationship between real commodity prices and the real exchange rate. Section III explains the sources and construction of the national real exchange rate and real commodity export price data used in this study. Section IV applies cointegration and vector error correction methodology to examine both the long-run and short-run determinants of the real exchange rate in commodity-dependent countries, especially the relationship between the real exchange rate and real commodity prices. It then questions whether the nominal exchange rate regime matters for commodity currencies; examines the relative volatility of the real exchange rate and real commodity prices of countries with commodity currencies; draws inferences regarding causality between the two series; and examines the speed of reversion of the real exchange rates to their commodity-price-dependent equilibria. Section V concludes.

II. Theoretical Framework

We consider a small open economy that produces two different types of goods: a nontradable good and an exportable good. For the purpose of our work, we associate the production of this exportable good with the production of a primary commodity (agricultural or mineral product). Nevertheless, our analysis is in line with the literature that stresses the role of the terms of trade in the determination of the real exchange rate, which includes (among others) work by De Gregorio and Wolf (1994) and Obstfeld and Rogoff (1996).

In our analysis, factors are mobile and the exportable good (as well as nontraded good) is *produced* domestically. Therefore, we abstract from demand-side considerations and concentrate on a representation of long-run relative price determination.[7] The details of the model are as follows.

A. Domestic Production

There are two different sectors in the domestic economy: one sector produces an exportable called "primary commodity"; the other sector consists of a continuum of firms producing a nontradable good. For simplicity, we assume that the production of these two different types of goods requires labor as the only factor. In particular, the production function for the primary commodity is:

$$y_X = a_X L_X \tag{1}$$

where L_X is the amount of the labor input demanded by the commodity sector and a_X measures how productive labor is in this sector. In a similar fashion, the nontraded good is produced through the production function:

$$y_N = a_N L_N \tag{2}$$

where a_N captures the productivity of labor in the production of this good and L_N is the employment of labor in the nontradable sector. Crucially, we assume that labor can move freely across sectors in such a way that labor wages must be the same across sectors. Profit maximization in both sectors yields the familiar conditions:

$$\begin{aligned} P_X &= \frac{w}{a_X} \\ P_N &= \frac{w}{a_N} \end{aligned} \tag{3}$$

[7] In a demand-side analysis, the direction of the effects of terms of trade shocks on the real exchange rate is indeterminate, as it depends on the relative magnitude of income and substitution effects. The traditional explanation for the link between real exchange rates and the terms of trade highlights the dominance of the income effect—a deterioration in the terms of trade (say through a rise in the price of importables) implies lower real national income, decreasing the demand for both imported (tradable) goods and nontraded goods in the commodity-exporting country. In equilibrium the relative price of nontradables will fall, yielding a real exchange rate depreciation (Diaz-Alejandro (1982)). However, there is also a substitution effect of terms of trade shocks. If countries can readily switch consumption between importables and nontradables in response to an adverse terms of trade shock, then the price of nontradables will rise, yielding a real appreciation of the exchange rate (Edwards (1989); Cashin and McDermott (2003)).

In equilibrium, the marginal productivity of labor must equal the real wage in each sector. We assume that the price of the primary commodity is exogenous for (competitive) firms in the commodity sector, and that there is perfect competition in the nontraded sector. Therefore, we can rewrite the price of the nontraded good in order to express it as a function of the price of the exportable and the relative productivities between the export and nontradable sectors. We obtain:

$$P_N = \frac{a_X}{a_N} P_X \tag{4}$$

Thus, the relative price of the nontraded good (P_N) with respect to the primary commodity (P_X) is completely determined by technological factors and is independent of demand conditions. Notice that an increase in the price of the primary commodity will increase the wage in that sector (see equation (3)). Given our freely mobile labor assumption, wages and prices will also rise in the nontraded sector.

B. Domestic Consumers

The economy is inhabited by a continuum of identical individuals that supply labor inelastically (with $L = L_X + L_N$) and consume a nontraded good and a tradable good. This tradable good is imported from the rest of the world and is not produced domestically. Our assumptions on preferences imply that the primary commodity is also not consumed domestically. Each individual chooses the consumption of the nontraded and tradable good to maximize utility, which is assumed to be increasing in the level of aggregate consumption given by:

$$C = \kappa C_N^{\gamma} C_T^{1-\gamma} \tag{5}$$

where C_N represents purchases of the nontraded good, C_T purchases of the imported good and $\kappa = 1/[\gamma^{\gamma}(1-\gamma)^{(1-\gamma)}]$ is an irrelevant constant. The minimum cost of one unit of consumption C is given by:

$$P = P_N^{\gamma} P_T^{1-\gamma} \tag{6}$$

where P_T is the price in local currency of one unit of the tradable good. As usual, P is defined as the consumer price index. Now, the law of one price is assumed to hold for the imported good:

$$P_T = \frac{P_T^*}{E} \tag{7}$$

where E is the nominal exchange rate, defined as the amount of foreign currency per local currency, and P_T^* is the price of the tradable (imported) good in terms of foreign currency. We now specify in more detail the rest of the world.

C. Foreign Production and Consumption

So far we have assumed that the primary commodity is not consumed by domestic agents and is therefore completely exported. In addition, the domestic economy also imports a good that is produced only by foreign firms.[8] The foreign region consists of three different sectors: a nontraded sector; an intermediate sector; and a final good sector. The nontraded sector produces a good that is consumed only by foreigners using labor as the only factor. The technology available for the production of this good is given by:

$$Y_N^* = a_N^* L_N^* \qquad \textbf{(8)}$$

The foreign economy also produces an intermediate good that is used in the production of the final good. This intermediate good is produced using labor as the only factor. In particular, the production function available to firms in this sector is represented by:

$$Y_I^* = a_I^* L_I^* \qquad \textbf{(9)}$$

Labor mobility across (foreign) sectors ensures that the (foreign) wage is equated across sectors.[9] Again, we can express the price of the foreign nontraded good as a function of relative productivities and the price of the foreign intermediate good:

$$P_N^* = \frac{a_I^*}{a_N^*} P_I^* \qquad \textbf{(10)}$$

The production of the final good involves two intermediate inputs. The first is the primary commodity (produced by several countries, among them our domestic economy). The second is an intermediate good produced in the rest of the world. Producers of this final good, also called the tradable good, produce it by assembling the foreign intermediate input (Y_I) and the foreign primary commodity (Y_X) through the following technology:

$$Y_T^* = \nu (Y_I^*)^{\beta} (Y_X^*)^{1-\beta} \qquad \textbf{(11)}$$

[8] When we refer to the foreign economy, we do not mean the rest of the world. The rest of the world also includes other countries producing the primary commodity.

[9] We assume that labor can freely move across sectors within each region (domestic and foreign) but cannot move across regions.

Now, it is straightforward to show that the cost of one unit of the tradable good in terms of the foreign currency is given by:

$$P_T^* = \left(P_I^*\right)^{\beta}\left(P_X^*\right)^{1-\beta} \tag{12}$$

Foreign consumers are assumed to consume the foreign nontraded good and this final good in the same fashion as the domestic consumers. They also supply labor inelastically to the different sectors. Therefore, the consumer price index for the foreign economy can be represented by:

$$P^* = \left(P_N^*\right)^{\gamma}\left(P_T^*\right)^{1-\gamma} \tag{13}$$

D. Real Exchange Rate Determination

It is now straightforward to show how the real exchange rate is determined in the domestic economy. First, we define the real exchange rate as the domestic price of the basket of consumption relative to the foreign price of a common basket of consumption (EP/P^*). Using equations (6) and (13) we can show that:

$$\frac{EP}{P^*} = \left(\frac{a_X}{a_I^*}\frac{a_N^*}{a_N}\frac{P_X^*}{P_I^*}\right)^{\gamma} \tag{14}$$

where the term P_X^*/P_I^* corresponds to the commodity terms of trade (or the price of the primary commodity with respect to the intermediate foreign good) measured in foreign prices, a_X/a_I^* reflects the productivity differentials between the export and import (foreign) sectors, and a_N^*/a_N accounts for the productivity differentials between the local and foreign nontraded sectors. These last two terms embody the Balassa-Samuelson effect—an increase in productivity in the commodity sector will tend to increase wages, which translates into an increase in the price of the nontraded good. As the relative price of the primary commodity is exogenously determined, the final effect will be an appreciation of the real exchange rate.

In the empirical analysis of this paper, we will be centering our work on explaining the evolution of the real exchange rate of commodity-dependent economies. That is, economies in which one of the major sources driving movements in the real exchange rate is fluctuations in the commodity terms of trade. How do fluctuations in the relative commodity price translate into movements in the real exchange rate? In our simple model, an increase in the international price of the primary commodity will increase wages in the commodity sector. As wages are equal across sectors, the increase in wages will raise the relative price of the nontraded good and, therefore, appreciate the real exchange rate.

III. Data

The data used to examine whether there is a relationship between the real exchange rate of individual countries and the real price of their commodity exports are monthly time series, obtained from the International Monetary Fund's *International Financial Statistics* (IFS) and *Information Notice System* (INS) databases over the period January 1980 to March 2002, which gives a total of 267 observations.

A. Real Exchange Rate

The definition of the real exchange rate is the real effective exchange rate (REER) based on consumer prices (line *rec*). As such, we will examine the behavior of REER based on: (i) the nominal effective exchange rate (NEER), which is the trade-weighted average of bilateral exchange rates vis-à-vis trading partners' currencies; adjusted for (ii) differentials between the domestic price level (which is the consumer price index) and the foreign price level (which is the trade-weighted average of trading partners' consumer price indices) (RELP). We analyze effective rather than bilateral real exchange rates as the effective rate measures the international competitiveness of a country against all its trade partners, and helps to avoid potential biases associated with the choice of base country in bilateral real exchange rate analyses.

The REER indices measure how nominal effective exchange rates, adjusted for price differentials between the home country and its trading partners, have moved over a period of time. The CPI-based REER indicator is calculated as a weighted geometric average of the level of consumer prices in the home country relative to that of its trading partners, expressed in a common currency. The International Monetary Fund's seasonally-adjusted, CPI-based REER indicator of country i is defined as: $REER_i = \left[(P_i R_i) / \exp \sum_{i=1}^{n} (W_{ij} \ln(P_j R_j)) \right]$, where j is an index that runs (from 1 to n) over country i's trade partner (or competitor) countries; W_{ij} is the trade weight attached by country i to country j, which are based on 1988-90 average data on the composition of trade in manufacturing, non-oil primary commodities and tourism services; P_i and P_j are the seasonally adjusted consumer price indices in countries i and j; and R_i and R_j are the nominal exchange rates of countries i and j's currencies in U.S. dollars.[10] The national REER series are expressed in logarithmic form (see Appendix I for additional details).

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[10] A decline (depreciation) in a country's REER index indicates a rise in its international competitiveness (defined as the relative price of domestic tradable goods in terms of foreign tradables). For a detailed explanation and critique of how the Fund's REER indices are constructed, see Zanello and Desruelle (1997) and Wickham (1993). As shown by McDermott (1996), alternative measures of the real exchange rate, such as real bilateral exchange rates based on consumer prices, and the IMF's REER based on normalized unit labor costs, are both highly correlated with the IMF's CPI-based REER index.

B. Real Commodity Price

The definition of the real price of commodity exports (RCOMP) is: the nominal price of commodity exports (NCOMP) deflated by the International Monetary Fund's index of (the unit value of) manufactured exports (MUV).[11] This paper follows Deaton and Miller (1996) and constructs NCOMP as a geometrically-weighted index of the nominal prices of 44 individual commodity exports, where for each country:

$NCOMP = \exp\left\{\sum_{k=1}^{K}(W_k\ (\ln P_k))\right\}$, *where* $W_k = ((P_{jk}\ Q_{jk})/(\sum_k P_{jk}\ Q_{jk}))$, P_k is the index of the dollar world price of commodity k (taken from the International Monetary Fund's *IFS*); W_k is the weighting item, which is the value of exports of commodity k in the total value of all K commodity exports, for the constant base period j; and Q is the quantity of exports of commodity k (taken from UN COMTRADE data).[12] Importantly, each country's NCOMP will be unique, because W_k is country specific.[13] The national RCOMP series are expressed in logarithmic form (see Appendix II for additional details).

Most previous studies of the macroeconomic effects of commodity price movements in developing countries have used either the prices of individual primary commodities (Cuddington and Urzua (1989)), terms of trade indices (Cashin and Pattillo (2000)) or aggregate (noncountry-specific) indices of commodity price movements (Grilli and Yang (1988)). The exceptions have been the country-specific indices of prices of commodity exports constructed by Deaton and Miller (1996) and Dehn (2000).[14] Few exporters of

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[11] This real price is also described in the literature as the commodity terms of trade. The manufactured unit value (MUV) index is a unit value index of exports from 20 industrial countries, and use of the MUV index as a deflator is common to most studies in the commodity-price literature (see Grilli and Yang (1988), Deaton and Miller (1996), Cashin, Liang and McDermott (2000)).

[12] In this paper, 'commodity exports' are defined as nonfuel primary product (agricultural and mineral primary products) exports—see Appendices I and II for additional details. In constructing national measures of the terms of trade based upon world prices, it is assumed that the country's export basket is predominantly composed of commodities, and its import basket is predominantly composed of manufactures. Given that the country is a price taker on world markets, the ratio of the world price of commodities to the world price of manufactures will be a close approximation to the terms of trade.

[13] Baxter and Kouparitsas (2000) and Kouparitsas (1997) show that, for nonfuel commodity exporters, the terms of trade is essentially the relative prices of their commodity exports and manufactured imports. Across both developing and developed countries, there is little variation in the import share devoted to manufactured goods (averaging about 65 percent of the import basket), nonfuel goods (20 percent), and fuels (15 percent). Accordingly, they find that cross-country differences in movements in the terms of trade largely emerge on the export price side.

[14] Our national commodity price indices differ from those of Deaton and Miller (1996) and Dehn (2000) as they are based on monthly, rather than quarterly or annual data, and cover an expanded range of individual commodities.

nonfuel commodities are so specialized that the export prices of a single commodity can well approximate movements in an index of commodity export prices based on the export baskets of individual commodity-exporting countries. In addition, terms of trade indices are too broad, as they also cover movements in noncommodity prices (Dehn (2000)). Terms of trade indices are also typically calculated using export and unit values, which are affected by the composition of exports and so by the composition of GDP (Deaton and Miller (1996)). Finally, movements in aggregate commodity price indices are likely to poorly represent the movements in country-specific commodity export price indices, as prices of individual commodities do not tend to move together on world commodity markets (Cashin, McDermott, and Scott (1999)).

Following Deaton and Miller (1996), the commodity export weights used in the construction of our national commodity price indices are held fixed over time as we are interested in constructing an *exogenous* variable, and so exclude volume effects of changes in commodity export prices.[15] In addition, one of the chief advantages of using world commodity prices is that they are typically exogenous to the behavior of individual countries. The exogeneity of world commodity prices is consistent with the small share of world commodity markets held by most developing countries, even for those commodities in which they are highly concentrated. Previous empirical analyses have concluded that commodity-exporting countries are price takers on world commodity markets, and have negligible long-term market power in the markets for their commodity exports (Mendoza (1995), Broda (2002), World Bank (1994)).

C. Potential Commodity-Currency Countries

In selecting commodity-dependent developing countries to be included in our sample, we followed the classification of developing countries used in the International Monetary Fund's *World Economic Outlook*, for the years 1988-92, the mid-point of our sample (IMF (1996a)). The International Monetary Fund classifies developing countries by the composition of their export earnings and other income from abroad, and has five categories: *fuel* (Standard International Trade Classification (SITC 3); *manufactures* (SITC 5 to 8, less 68); *nonfuel primary products* (SITC 0, 1, 2, 4, and 68); *services, income, and private transfers* (exporters of services and recipients of income from abroad, including workers' remittances); and *diversified export earnings*. Countries whose 1988-92 export earnings in any of the first four categories accounted for more than half of total export earnings are allocated to that group, while countries whose export earnings were not dominated by any of the first four categories are defined as countries with diversified export earnings (see IMF (1996a) and Appendix III).

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[15] Deaton and Miller (1996) point out that while a geometric weighted average index uses fixed base-year weights (and so cannot capture changes in the structure of trade occurring outside the base period), an advantage is that by holding volumes constant, the index is driven by price rather than quantity movements. However, the indices will tend to downplay the wealth effects of commodity price movements, as they do not account for volume effects of such price changes.

Those developing countries in the IMF's category of *nonfuel primary products* are included in our sample, as are those in the category *diversified export earnings*, as many of these countries derive a large (yet not dominant) share of their export earnings from the export of nonfuel primary products. On this basis, the number of countries with potential commodity currencies is 73. Of these 73 countries, 12 were excluded from our analysis due to the unavailability of a consistent time series of data on their real effective exchange rate, leaving 61 developing countries in our sample. Of these 61 countries, 8 were excluded due to the unavailability of UN COMTRADE data on their commodity exports, leaving 53 developing countries in our sample. In addition, 5 commodity-dependent *industrial countries* (Australia, Canada, Iceland, Norway, and New Zealand) were included in our sample, to compare and contrast their results with those of the commodity-dependent developing countries. The 58 countries in our sample are listed (by geographic region) in Appendix III.

As expected, the export of commodities is a major source of export income for the 58 countries in our sample of commodity-exporting countries. In Table 1 we report the export share of the three most important commodity exports, and the total export share of the 44 individual commodities used to construct the indices of the nominal world price of national export baskets. During the 1990s, the cross-country mean share of total export receipts derived from primary commodity exports was about 48 percent. Among sub-Saharan African countries, commodity exports typically exceeded 50 percent of total exports, especially for Burundi (97 percent), Madagascar (90 percent), and Zambia (88 percent). Even among developed countries, the share of primary commodity exports in total exports is quite high (Australia, 54 percent; Iceland 56 percent). In addition, many countries remain overwhelmingly dependent on export receipts from their dominant commodity exportable—cases where the dominant exportable exceeded 90 percent of commodity export receipts include Niger (uranium), Dominica (bananas), Ethiopia (coffee), Zambia (copper), and Mauritius (sugar) (see Table 1).

The REER data (base 1990=100) for all countries are set out in Figures 1 to 10—an increase in the REER series indicates a real appreciation of the country's currency. Several features of the data stand out. First, a cursory inspection of the REER series indicates that most countries have real exchange rates that appear to exhibit symptoms of drift or nonstationarity. There appear to be substantial and sustained deviations from purchasing power parity (that is, nonstationarity in the REER). Typically, the evolution of the REER appears to be a highly persistent, slow-moving process; for most countries the REER does not appear to cycle about any particular equilibrium value, especially for countries such as Ecuador and India (the general depreciation of its exchange rate is typical of a process with a unit root). Second, sharp movements in the REER during the 1980s and 1990s are a relatively frequent occurrence, especially for countries experiencing rapid nominal devaluations, such as the countries of the CFA franc zone (for example, Mali and Togo).

Table 1. Principal Commodity Exports and
Share of Primary Commodities in Total Exports, 1990–99

Country	Principal Exports			Share of Exports			
	1	2	3	1	2	3	44
Argentina	Soy Meal	Wheat	Maize	18	13	11	41
Australia	Coal	Gold	Aluminum	22	14	12	54
Bangladesh	Shrimp	Tea	Fish	76	15	8	8
Bolivia	Zinc	Tin	Gold	27	18	13	56
Brazil	Iron	Coffee	Aluminum	21	15	10	35
Burundi	Coffee	Gold	Tea	59	35	2	97
Cameroon	Cocoa	Hardwood Logs	Aluminum	23	22	14	53
Canada	Softwood Sawn	Aluminum	Wheat	28	14	12	16
Central African Republic	Cotton	Coffee	Softwood Logs	82	9	5	43
Chile	Copper	Fish	Fishmeal	70	9	6	58
Colombia	Coffee	Coal	Bananas	48	19	18	40
Costa Rica	Bananas	Coffee	Fish	43	33	5	31
Côte d'Ivoire	Cocoa	Coffee	Cotton	65	14	6	65
Dominica	Bananas	Tobacco		98	1		32
Ecuador	Bananas	Shrimp	Coffee	45	30	8	49
Ethiopia	Coffee	Hides	Cotton	91	5	2	71
Ghana	Cocoa	Gold	Aluminum	61	24	7	72
Guatemala	Coffee	Sugar	Bananas	47	24	14	49
Honduras	Coffee	Bananas	Shrimp	47	30	6	67
Iceland	Fish	Aluminum	Shrimp	73	20	7	56
India	Rice	Shrimp	Soy Meal	18	15	12	31
Indonesia	Crude Petroleum	Natural Gas	Natural Rubber	34	23	7	43
Kenya	Tea	Coffee	Fish	53	30	5	45
Madagascar	Coffee	Shrimp	Sugar	42	40	6	39
Malawi	Tobacco	Tea	Sugar	78	8	7	90
Malaysia	Palm Oil	Natural Rubber	Hardwood Logs	44	15	15	13
Mali	Cotton	Gold		88	12		85
Mauritania	Iron	Fish	Gold	65	34	1	64
Mauritius	Sugar	Wheat		97	1		27
Mexico	Crude Petroleum	Copper	Coffee	72	5	5	15
Morocco	Phosphate Rock	Fish	Lead	55	14	7	14
Mozambique	Cotton	Sugar	Maize	33	19	9	26
Myanmar	Hardwood Logs	Rice	Shrimp	60	18	7	52
New Zealand	Lamb	Beef	Wool	20	17	14	36
Nicaragua	Coffee	Beef	Shrimp	32	15	14	69
Niger	Uranium	Tobacco		96	3		68
Norway	Crude Petroleum	Natural Gas	Fish	67	13	8	63
Pakistan	Rice	Cotton	Sugar	46	28	13	12
Papua New Guinea	Copper	Gold	Palm Oil	23	23	20	59
Paraguay	Soybeans	Cotton	Soy Meal	44	26	9	79
Philippines	Coconut Oil	Copper	Bananas	29	21	12	10
Peru	Copper	Fishmeal	Gold	28	19	15	69

Table 1 (Concluded). Principal Commodity Exports and
Share of Primary Commodities in Total Exports, 1990–99

Country	Principal Exports 1	Principal Exports 2	Principal Exports 3	Share of Exports 1	Share of Exports 2	Share of Exports 3	Share of Exports 44
Sri Lanka	Tea	Natural Rubber	Tobacco	78	9	6	20
St. Vincent & Grenadines	Bananas	Wheat	Rice	60	23	17	72
Sudan	Cotton	Gold	Sugar	45	12	12	44
Suriname	Aluminum	Rice	Nickel	80	8	5	86
Syrian Arab Republic	Crude Petroleum	Cotton	Wheat	88	8	2	74
Tanzania	Coffee	Tobacco	Cotton	27	18	17	59
Thailand	Rice	Natural Rubber	Shrimp	26	24	23	16
Togo	Phosphate Rock	Cotton	Coffee	44	40	9	84
Tunisia	Tobacco	Phosphate Rock	Shrimp	23	21	20	8
Turkey	Tobacco	Wheat	Sugar	34	16	14	8
Uganda	Coffee	Fish	Gold	71	8	4	84
Uruguay	Beef	Rice	Fish	36	27	13	32
Zambia	Copper	Sugar		97	2		88
Zimbabwe	Tobacco	Cotton	Nickel	58	8	8	54

Sources: United Nations (COMTRADE); International Monetary Fund, commodity price indices.

Notes: Columns marked 1–3 denote the three largest commodity exports of each country, and their share of total commodity exports. The column marked 44 denotes the share of total exports of goods that the 44 commodities tracked by the IMF comprise, and which were used in the construction of the nominal national commodity-price series.

Figures 1 to 10 also display the RCOMP indices (base 1990=100) for the countries in our sample. Using ocular regression methods, it appears that many countries display a close relationship between their real commodity prices and real exchange rates (such as Australia and Papua New Guinea), while others appear to display a close relationship once a one-time real effective depreciation is accounted for (such as several of the CFA franc zone countries of sub-Saharan Africa).

IV. Empirical Analysis of Comovement

We use the Engle and Granger (1987) cointegration approach to assess whether the level of real exchange rates and real commodity prices move together over time. That is, we examine whether there is a long-run relationship between real exchange rates and real commodity prices, which implies that deviations from any long-run relationship are self-correcting. For those countries where cointegration can be established between real exchange rates and real commodity prices, we then ascertain the direction of causality between the two series using the vector error correction methodology of Engle and Granger (1987). Finally, we measure the speed with which the real exchange rate of 'commodity currencies' revert to both their constant equilibrium level (as implied by PPP) and their time-varying equilibrium level (as implied by commodity-price-augmented PPP).

Figure 1. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3

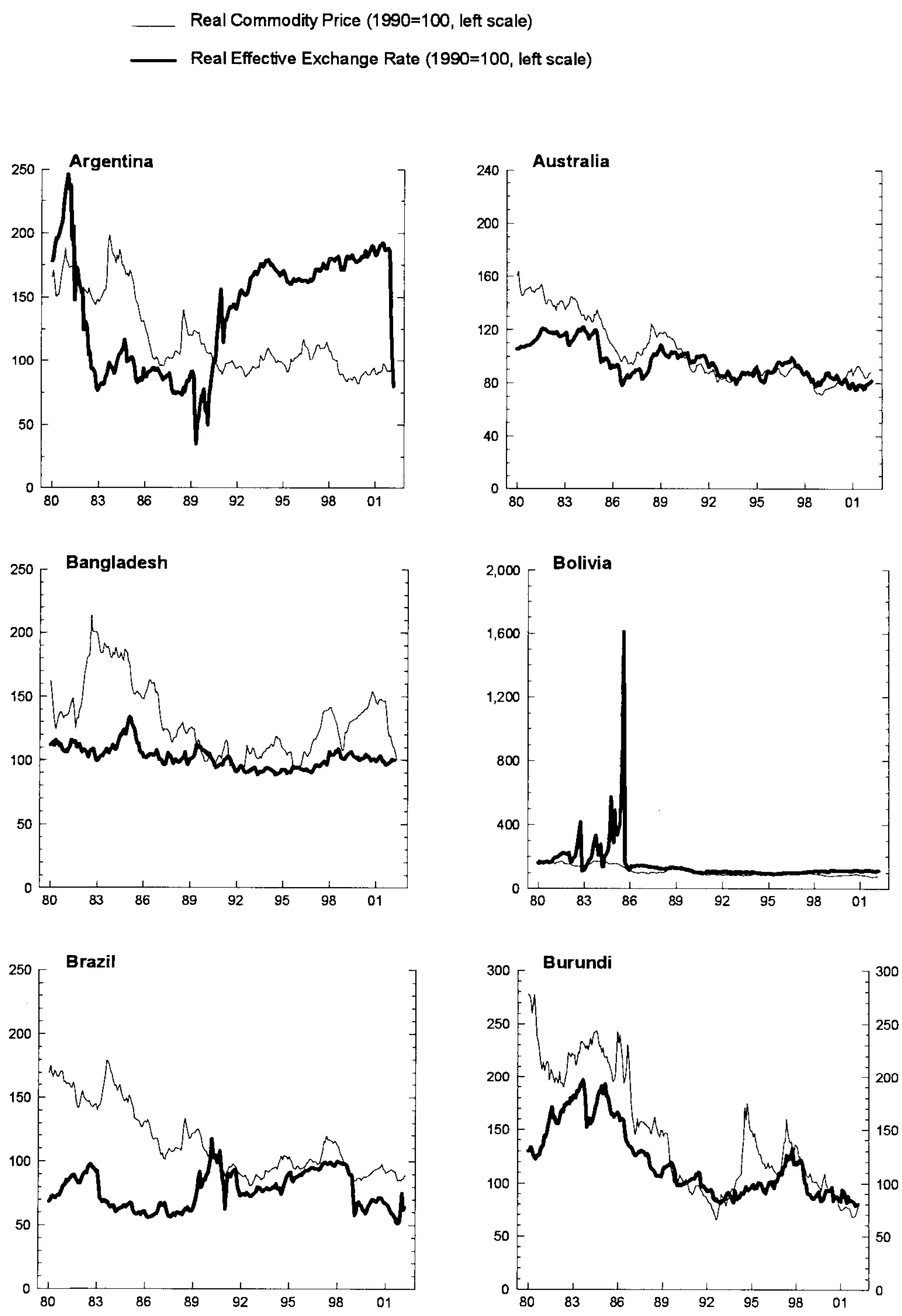


Figure 2. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3

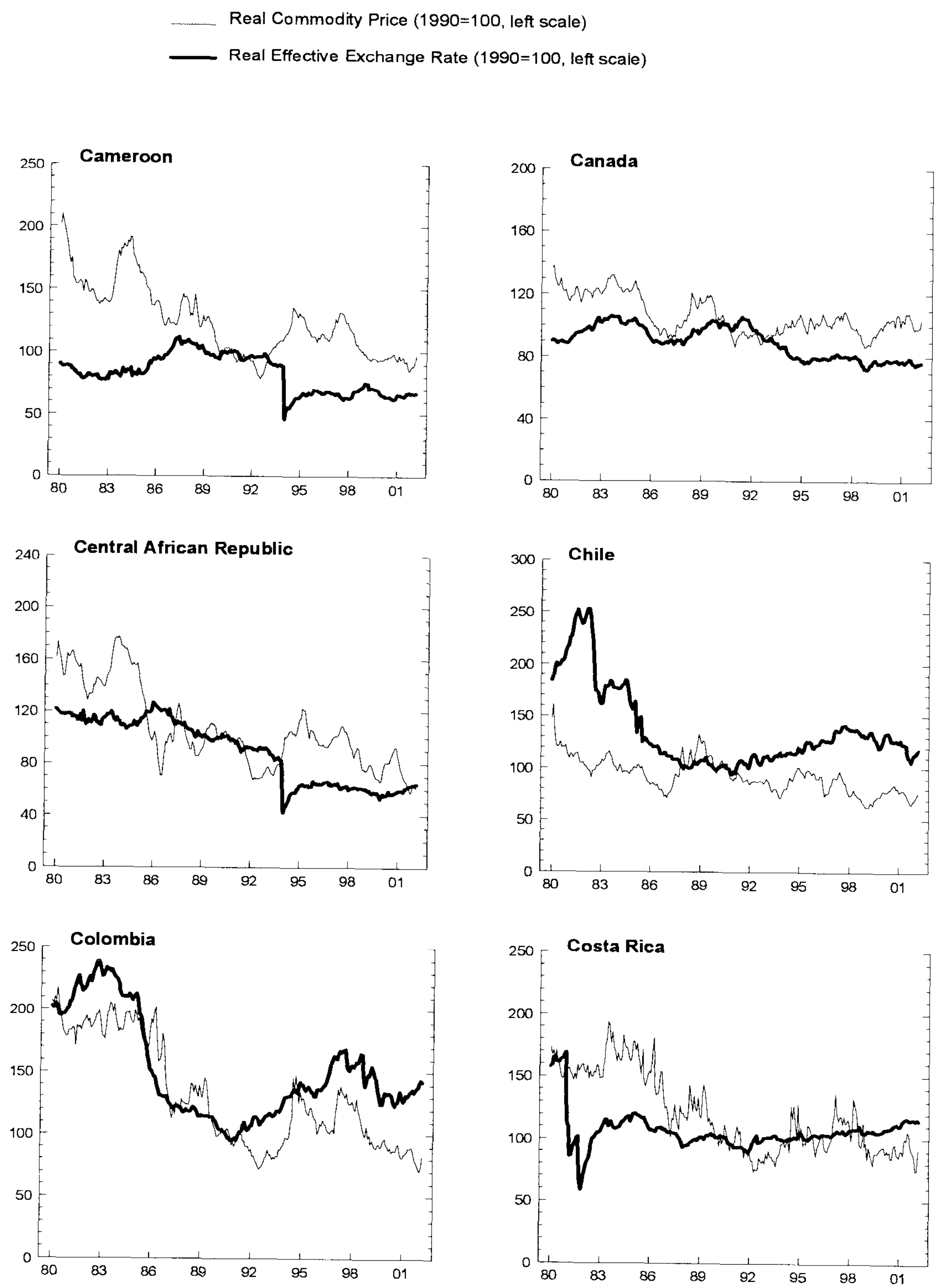


Figure 3. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3

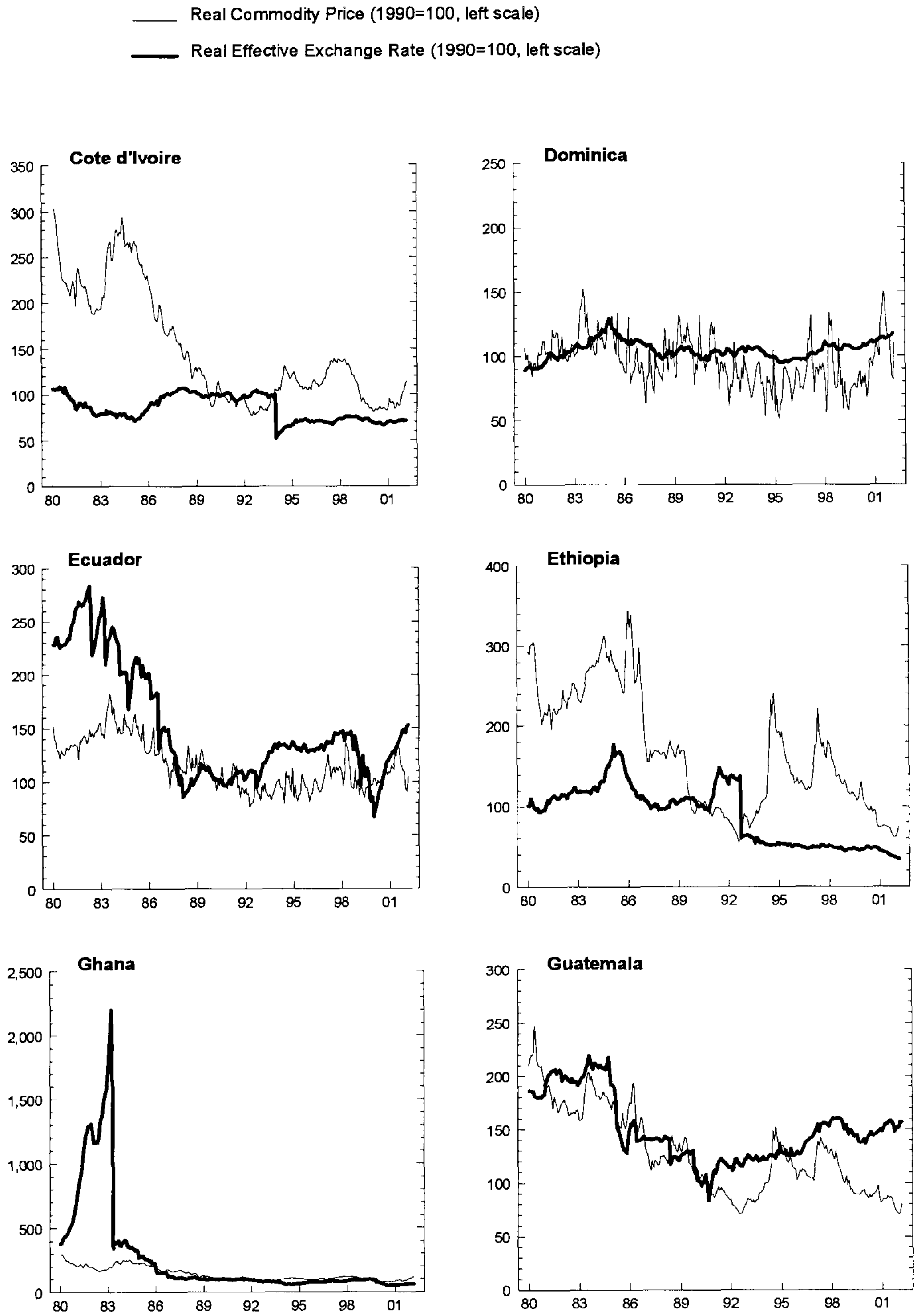


Figure 4. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3

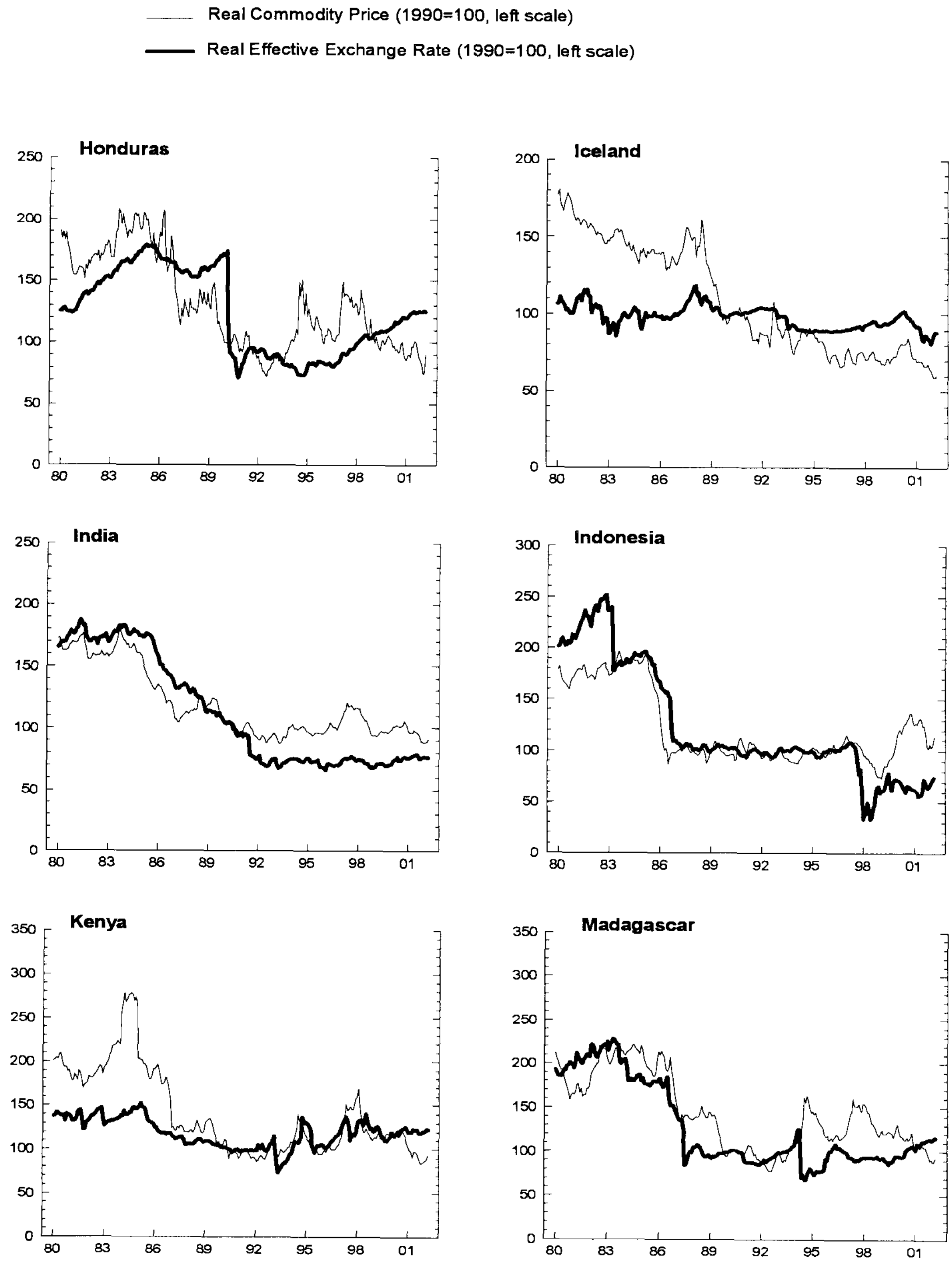


Figure 5. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3

Real Commodity Price (1990=100, left scale)

Real Effective Exchange Rate (1990=100, left scale)

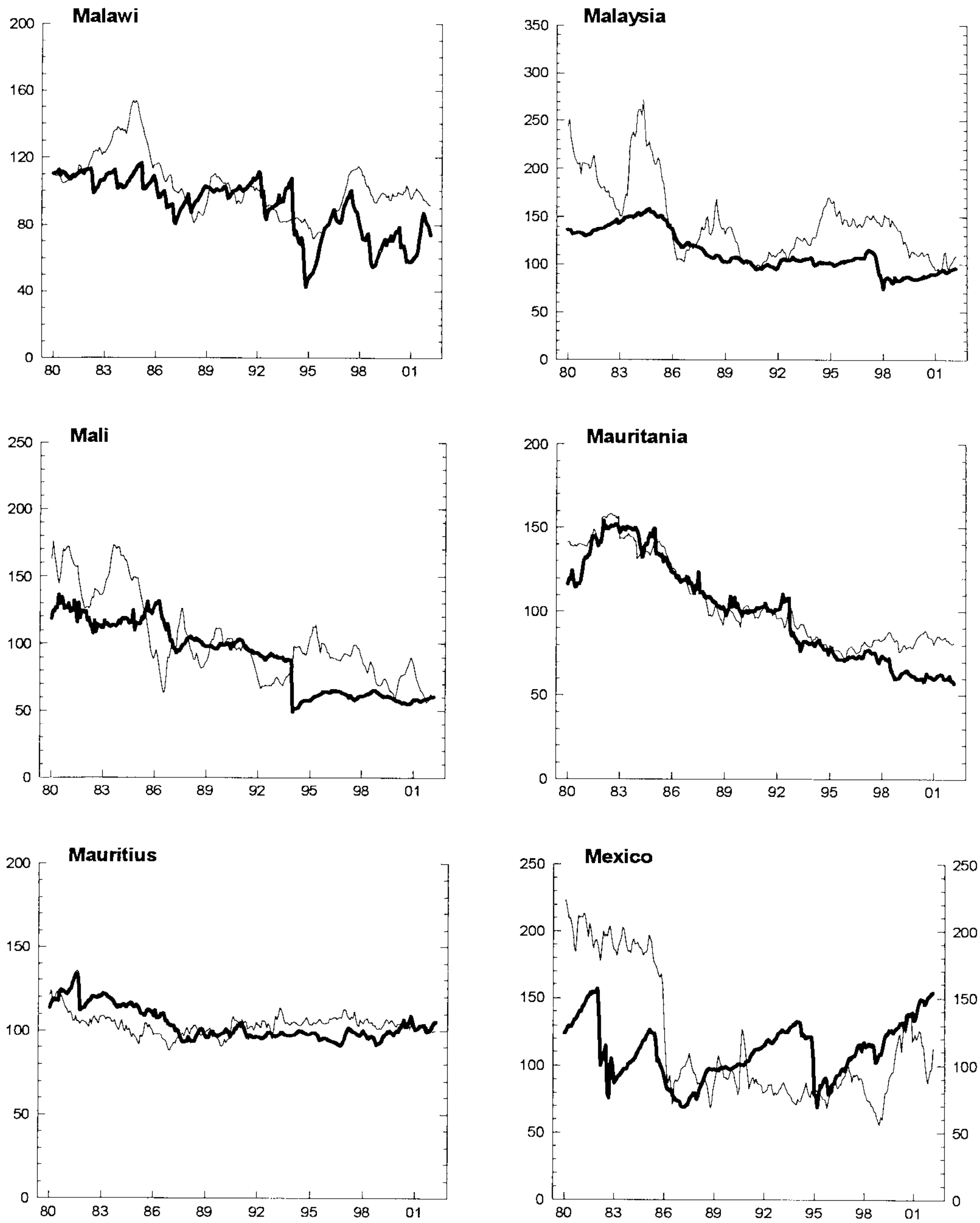


Figure 6. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3

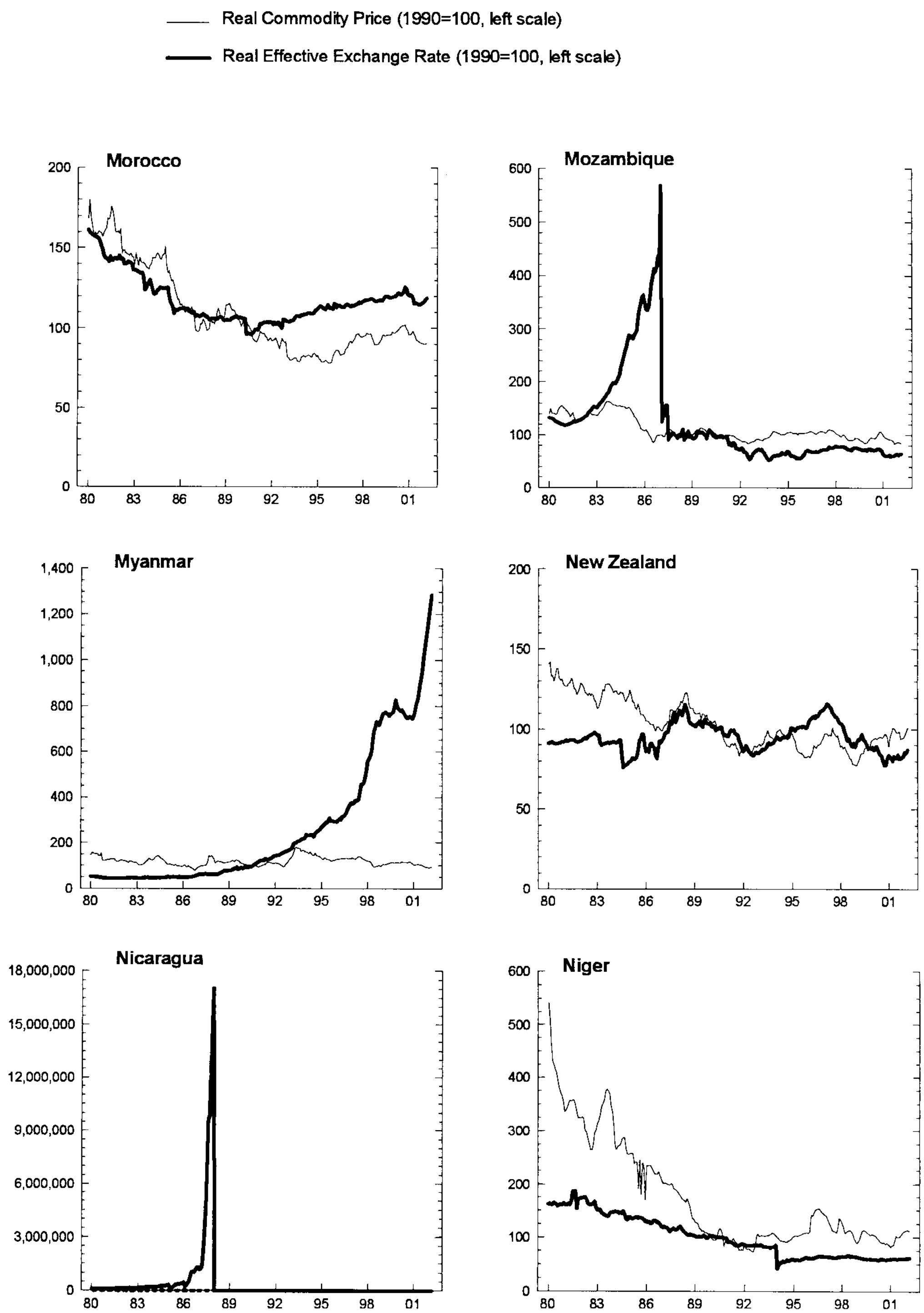


Figure 7. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3

Real Commodity Price (1990=100, left scale)

Real Effective Exchange Rate (1990=100, left scale)

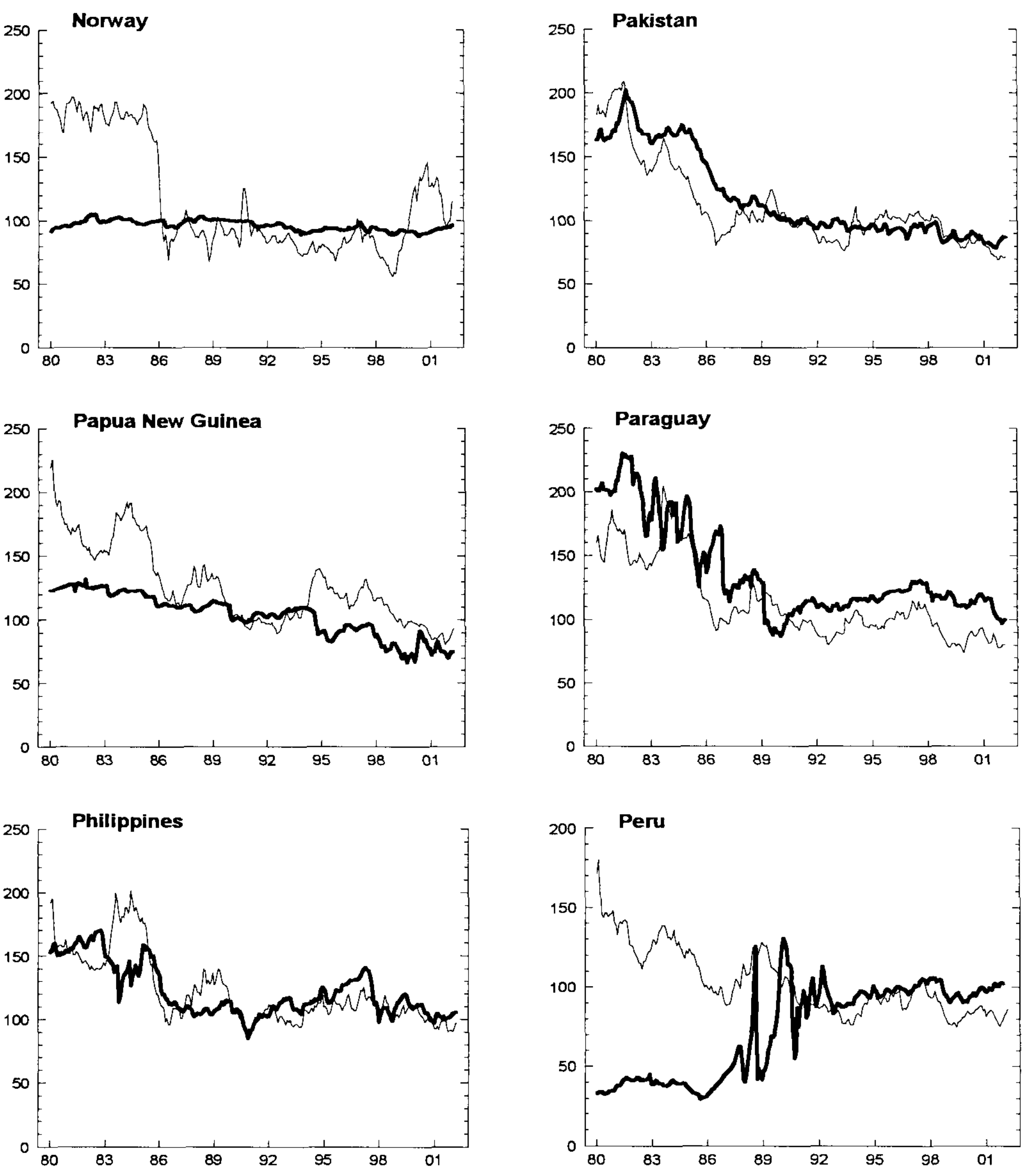


Figure 8. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3

——— Real Commodity Price (1990=100, left scale)

——— Real Effective Exchange Rate (1990=100, left scale)

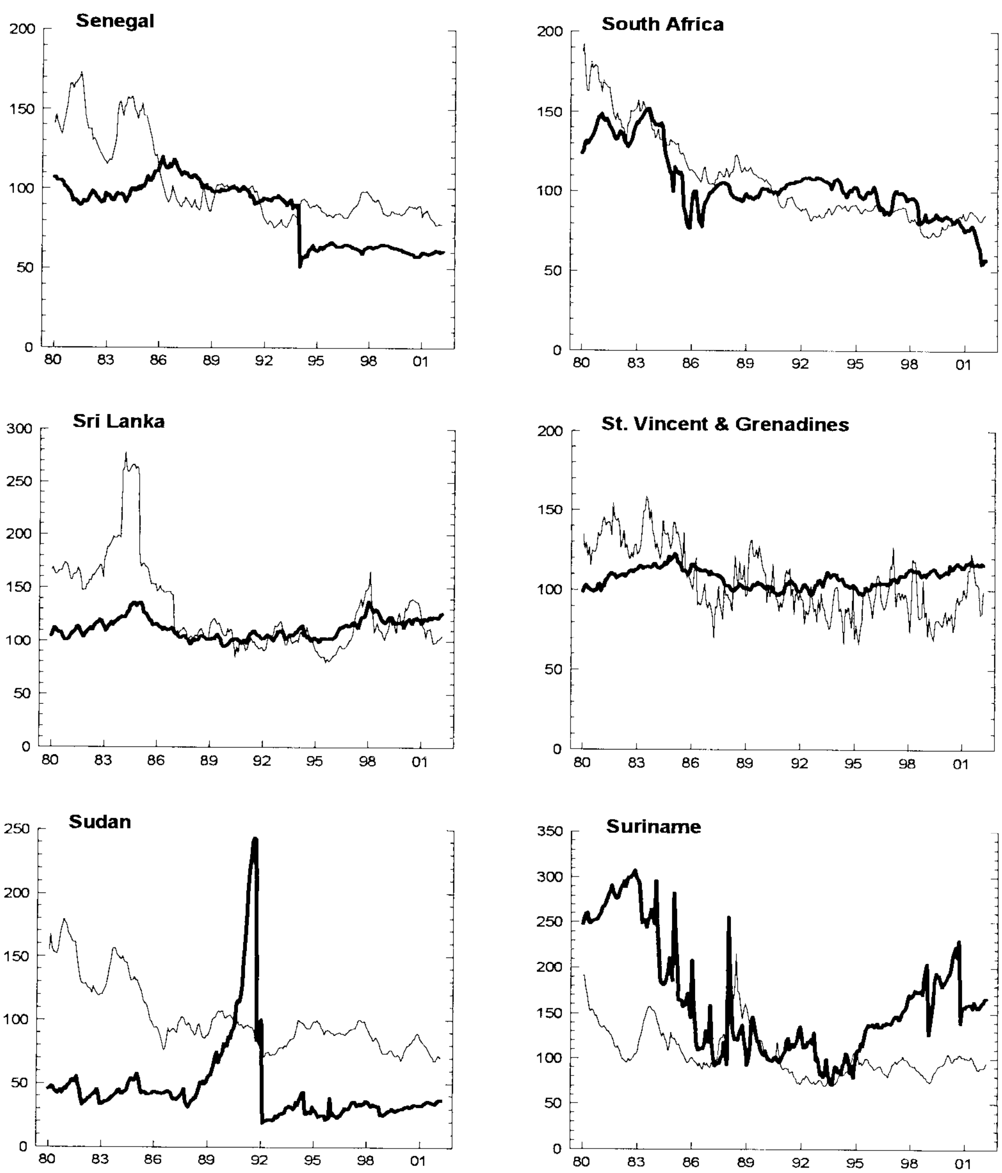


Figure 9. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3

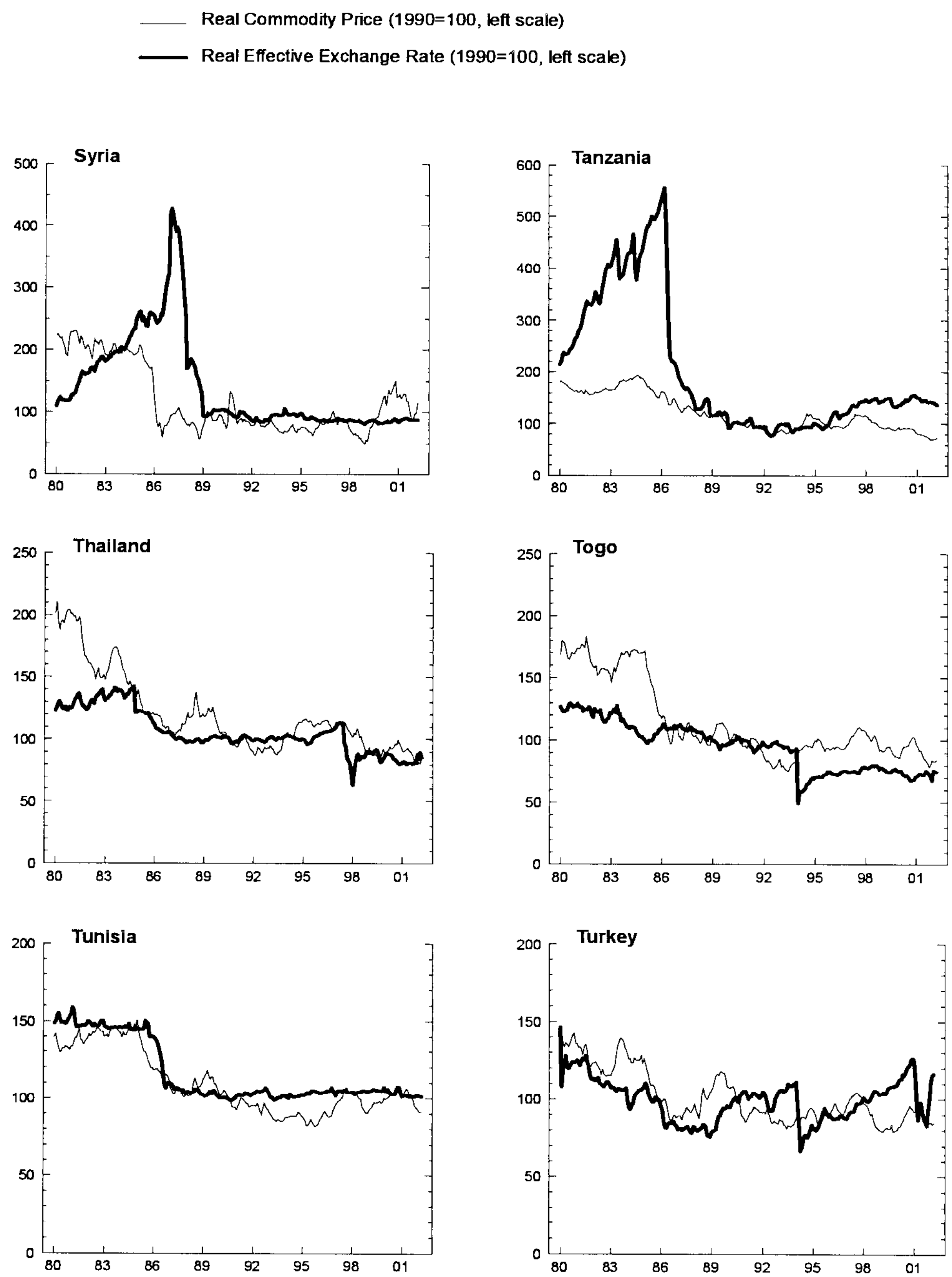
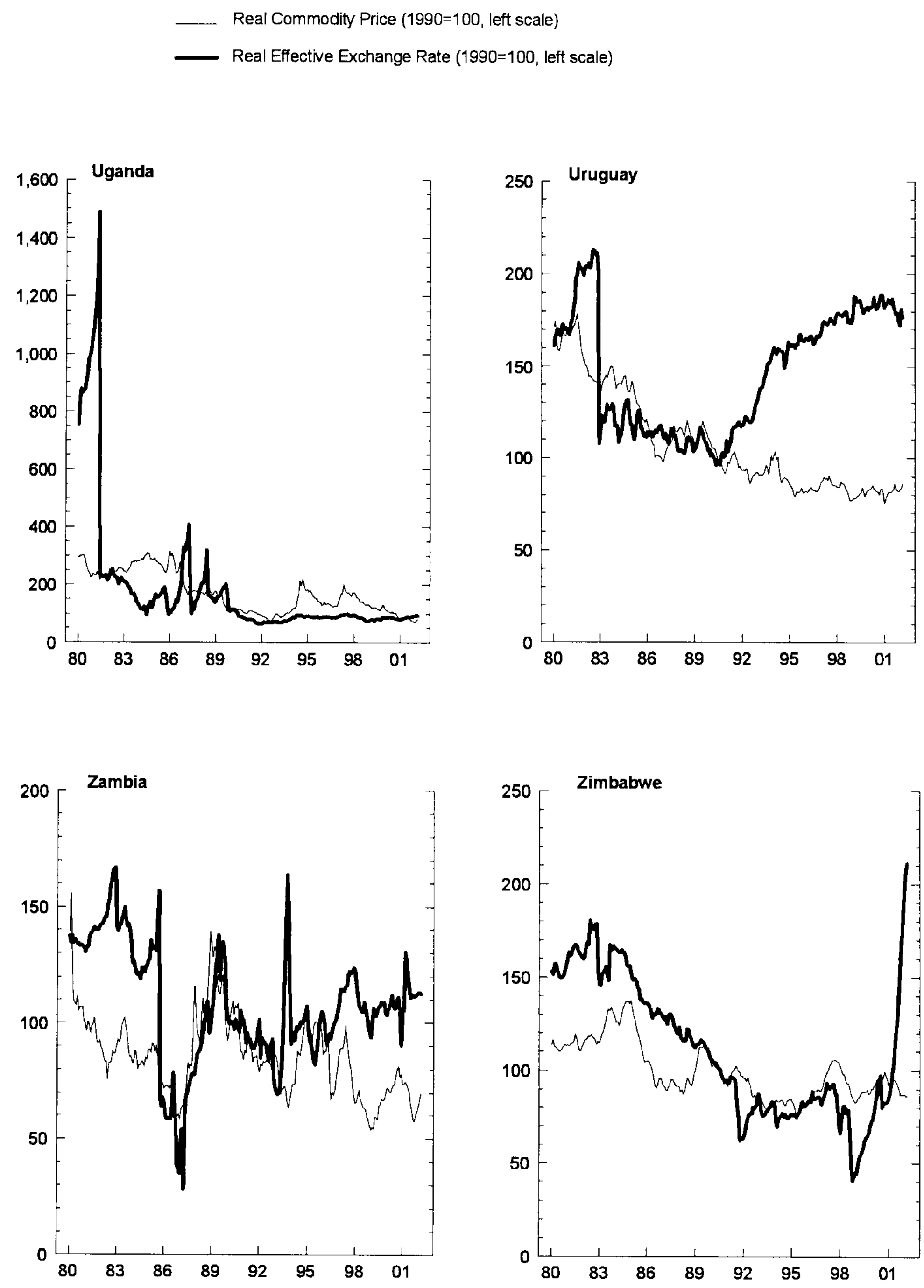


Figure 10. Real Exchange Rate and Real Commodity Price, Commodity-Exporting Countries, 1980:1 - 2002:3



A. Is There A Long-Run Relationship Between National Real Exchange Rates and Real Commodity Prices?

Economic theory has established that the long-run (equilibrium) real exchange rate is determined by the long-run value of certain 'fundamentals', such as the terms of trade, real interest rate differentials, and productivity differentials. Deviations of the actual real exchange rate from the equilibrium real exchange rate dictated by these fundamentals should be short-lived. If the fundamental determinants of the real exchange rate are integrated processes, it follows that the real exchange rate itself must also be an integrated process, and we can examine whether there is a long-run (cointegrated) relationship between the fundamentals and the real exchange rate.

As set out in Section II of the paper, for commodity-dependent countries the fundamental determinant of their real exchange rate are real commodity prices. In conducting our analysis we test, for each country, several hypotheses. First, that its REER and RCOMP series are nonstationary. Second, whether for each country there is a long-run (cointegrating) relationship between its real exchange rate and the real price of its commodity exports. Third, given that we establish cointegration, we test for parameter instability in the cointegrated model.

Order of Integration of the Series

We use the Phillips-Perron (1988) and Kwiatkowski et al. (1992) unit root tests to assess the time-series properties of our data. While the Phillips-Perron test maintains the null hypothesis of nonstationarity of the time series, the Kwiatkowski test uses a null hypothesis of stationarity. For both tests we include a constant term and trend in the fitted regression, and we employ the Bartlett kernel with Andrews' (1991) automatic bandwidth selector and the pre-whitened kernel estimator of Andrews and Monahan (1992). The results for both tests give very little evidence for stationarity—they indicate that for all countries both series (REER and RCOMP) were typically nonstationary in levels and stationary in first difference form.[16] The results of these tests for the stationarity of the real exchange rate are consistent with those of earlier work (see Boyd and Smith (1999)). Similarly, shocks to world commodity prices have been found to be highly persistent (Cashin, Liang, and McDermott (2000)).

[16] We also applied the Zivot-Andrews (1992) unit root test which allows for an exogenous change in the level of the series—with a few exceptions, all test statistics for the two series are again not statistically significant, indicating nonrejection of the unit root null. Accordingly, we conclude that the REER and RCOMP series of most countries exhibit behavior consistent with unit root nonstationarity in levels. Although not consistent with every test result (using these unit root tests there is some conflicting evidence as to whether the REER series of Mali, Mauritania, and Togo, and the RCOMP series of Iceland, are nonstationary in levels) these conclusions seem reasonable. The detailed results of the various unit root tests are available from the authors.

One possible reason for the failure to reject the null hypothesis of nonstationarity of the real exchange rate is that there may be macroeconomic disturbances, such as shocks to real commodity prices, which induce persistent deviations of real exchange rates from purchasing power parity. If the observed deviation from parity of each country's real exchange rate is caused by real commodity prices, then real exchange rates can be expected to be cointegrated with real commodity prices. Accordingly, in subsequent sections we treat real exchange rates and real commodity prices as I(1) variables, and go on to examine (for each country) whether there is a long-run relationship between these series for the period 1980-2002.

Examining for Cointegration: Allowing for Structural Change

Gregory and Hansen (1996a) demonstrate that the power of standard tests for cointegration falls when no allowance is made for structural shifts in the relationship between nonstationary series. Accordingly, the first step in the estimation procedure is to allow for the possibility that the cointegrated (long-run) relationship between the real effective exchange rate ($REER_t$) and real commodity price ($RCOMP_t$), has shifted at an unknown point in the sample. The possibility of a structural shift is allowed for because the period 1980-2002 has been marked by some significant policy framework changes in many countries, such as sharp nominal exchange rate adjustments and changes in nominal exchange rate regime, and by rapid fluctuations in the world prices of many primary commodities.[17] This period provides a very severe test of the commodity-currency model of real exchange rate movements, and suggests there is a possibility of a regime shift in behavior as economic agents adapt to any new economic environment. Moreover, the timing of any such regime shift is likely to be unknown, because there is not necessarily a one-to-one correspondence between potential causes of a regime shift and its occurrence in the data. Use of the Gregory-Hansen (1996a) test for cointegration is therefore helpful in this instance, since it allows for the timing of any regime shift to be unknown a priori.[18]

Gregory and Hansen (1996a) commence with the standard model for cointegration in the presence of no structural change, viz:

$$REER_t = \beta_0 + \beta_1 RCOMP_t + \varepsilon_t, \qquad t=1,...,T, \qquad \textbf{(15)}$$

[17] Goldfajn and Valdés (1999) find that real exchange rate overvaluations are typically corrected through sharp nominal devaluations (involving a collapse in the nominal exchange rate regime), rather than through cumulative inflation differentials.

[18] As argued by Gregory, Nason, and Watt (1996), it is likely that any failure to find cointegration between REER and the RCOMP may be due to the presence of a structural shift in the cointegrating relationship, which if present would bias standard residual-based tests of the null hypothesis of no cointegration towards not rejecting the null.

where REER and RCOMP are I(1) variables, and the residual ε_t is I(0). In the context of the data considered here, there is an apparent level shift in the long-run relationship between the real exchange rate and real commodity price series, which typically occurs as a level shift in the real (and nominal) exchange rate. Accordingly, as an alternative to equation (15), Gregory and Hansen propose a model where structural change occurs with a shift in the intercept term:

$$REER_t = \beta_0 + \beta_1 RCOMP_t + \beta_2 \varphi_{t\pi} + \varepsilon_t , \qquad t = 1,...,T, \quad \textbf{(16)}$$

where β_0 denotes the cointegrating intercept coefficients before the shift, β_2 denotes the change in the intercept coefficients, and RCOMP and ε_t are as described above. Importantly, structural change is modeled using the following dummy variable:

$$\varphi_{t\pi} = \begin{cases} 0 \ \ if \ \ t \leq [T\pi] \\ 1 \ \ if \ \ t > [T\pi], \end{cases} \quad \textbf{(17)}$$

where the unknown parameter $\pi \in (0,1)$ denotes the timing of the change point in terms of a fraction of the sample and [] denotes integer part. Given that the timing of shifts ($T\pi$) in the relationship between macroeconomic series is unlikely to be known a priori, the Gregory-Hansen test for shifts in cointegrated models is useful as it does not require information on the timing of the such events.

A test of the null hypothesis of no cointegration is run, against the alternative hypothesis given by equation (16). In doing so, the usual Phillips-Perron (1988) $Z(t)$ cointegration test statistic is computed for each possible shift $\pi \in \Pi$, using the residuals from the cointegrating regression of equation (16). The π is chosen so that $Z(t)$ takes the smallest value (largest negative value) across all possible break points, where Π is any compact subset of (0,1) since the smallest $Z(t)$ gives the least favorable result for the null hypothesis (that is, the greater chance of rejecting the null of no cointegration). We will denote the smallest of these $Z(t)$ statistics as $Z(t)^*$.[19]

While the Gregory-Hansen (1996a) test was designed to investigate if there is a cointegrating relation after allowing for a structural shift, the test also has power to detect cointegration when there is no structural shift. Consequently, a rejection of the null hypothesis of no cointegration may not be indicative of changes in the cointegrating vector, as the existence of a stable cointegrating relationship could also induce such a rejection. Accordingly, Gregory and Hansen (1996b) recommend that it is also necessary to test for cointegration using standard statistics that assume a stable cointegrating relation.

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[19] Following Gregory and Hansen (1996a), Π is here taken to be the compact subset $\Pi = [0.15T, 0.85T]$.

The Phillips-Ouliaris (1990) cointegration statistics test the null hypothesis of no cointegration between REER and RCOMP against the alternative hypothesis of a stable cointegration relationship. The null of the Gregory-Hansen (1996a) model is also no cointegration between $REER_t$ and $RCOMP_t$, while the alternative hypothesis is cointegration with a one-time structural shift of unknown timing in the cointegrating relationship (change in cointegrating intercept coefficients). Note that if the conventional cointegration test (such as the Phillips-Ouliaris $Z(t)$ and $Z(\alpha)$ tests) does not reject the null of no cointegration but the Gregory-Hansen $Z(t)^*$ test does, then there is evidence of a structural shift in the cointegrating relationship (Gregory and Hansen (1996a)).

The results of the Gregory-Hansen (1996a) cointegration test are set out in Appendix IV. For 19 countries, the Gregory-Hansen statistics are consistent with a long-run cointegrating relationship between REER and RCOMP (allowing for a structural shift), as conventional cointegration tests cannot reject the null of no cointegration but the Gregory-Hansen test does. Importantly, significant values of the test statistic appear to broadly coincide with periods of nominal exchange rate revaluation, such as the 1994 devaluation of the nominal exchange rate of the CFA franc zone countries (Reinhart and Rogoff (2002)). Finally, we find that for all but 10 of the 58 countries, the Phillips-Ouliaris $Z(t)$ and $Z(\alpha)$ statistics are too small to reject the null of no cointegration (see Appendix IV).

Importantly, if *both* conventional cointegration tests and the Gregory-Hansen test reject the null hypothesis of no cointegration (as occurs for Bolivia, Costa Rica, and Kenya), then while it is clear that there is strong evidence in favor of a long-run relationship, it is unclear whether a structural shift has occurred because (as noted above) the Gregory-Hansen test is powerful against conventional cointegration. In this case, further investigation is necessary to enable a distinction to be drawn between cointegration with stable parameters and cointegration with a structural shift, as the null hypothesis of no cointegration is rejected in comparison with either alternative hypothesis. Gregory and Hansen (1996b) suggest using Hansen's (1992) parameter instability tests (which are based on the residuals of a FM least squares regression), where the null hypothesis is cointegration with stable parameters, to determine whether there has been a shift in the cointegration relationship. For all three Hansen (1992) tests, the null hypothesis is that the cointegrating parameters are constant, while the alternative hypothesis is no cointegration due to a change in the parameters at some unknown point in the sample. In particular, under the alternative hypothesis of parameter instability, the SupF test is focused on any abrupt shift in the cointegrating vector; the MeanF and Lc tests detect any gradual changes in the regression coefficients.[20] Using the Hansen (1992) tests we find no evidence of unstable relationship between REER and RCOMP for any of the above three countries, and so conclude that there is cointegration with stable parameters.

[20] The MeanF and SupF tests require truncation of the sample of size T to avoid the test statistics diverging to infinity—we follow Hansen (1992) and use the subset $[0.15T, 0.85T]$.

B. Cointegration Results and Long-Run Elasticity Estimates

For those 19 countries where the null hypothesis of no cointegration could be rejected using the $Z(t)^*$ test, the cointegrating relationship between each country's REER and RCOMP (as set out in equation (16) above) was estimated using Phillips and Hansen's (1990) Fully Modified (FM) method. FM estimation is a semiparametric procedure that modifies least squares regression to account for potential endogeneity of the regressors and serial correlation caused by cointegrating relationships.[21] The FM method yields an asymptotically correct variance-covariance estimator when estimating cointegrating vectors in the presence of serial correlation and endogeneity—the results are set out in the lower panel of Table 2.[22, 23] Estimates of the commodity price elasticity of the real exchange rate are typically positive, while there is typically a downward shift in the constant term in the cointegrating regression. All cointegrating regressions have excellent explanatory ability, with coefficients of determination ranging between about 0.7 and 0.95. This is consistent with real commodity prices having a strong influence on movements in real exchange rates for those countries with commodity currencies.

For those ten countries where the null hypothesis of no cointegration could be rejected using the Phillips-Ouliaris $Z(t)$ or $Z(\alpha)$ tests, the cointegrating relationship between each country's REER and RCOMP (as set out in equation (15) above) was again estimated using Phillips and Hansen's (1990) FM method—the results are set out in the upper panel

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[21] For the Phillips-Hanson FM estimation we employ the Bartlett kernel, Andrews' (1991) automatic bandwidth selector and the pre-whitened kernel estimator of Andrews and Monahan (1992). The regression was run without a trend term, which was found to be not statistically significantly different from zero in the cointegrating regressions. This absence of a significant time trend in the cointegrating regressions indicates that, controlling for real commodity prices, there is little support for sectoral productivity differentials (the Balassa-Samuelson effect) driving commodity-currency real exchange rates.

[22] Ordinary least squares estimation could be used to yield consistent estimates of the cointegrating parameters. However, least squares estimation is inefficient and yields nonstandard distributions of the estimators, making standard inference tests problematic in the least squares framework, while these difficulties are overcome in the FM method (Phillips and Hansen (1990)). Importantly, FM-based estimates are robust to any potential endogeneity of real commodity prices.

[23] While the null hypothesis of no cointegration could be rejected in favor of the alternative hypothesis of cointegration (allowing for a structural shift) for Ethiopia, Madagascar, Mauritius, Peru, and Senegal, for these countries the coefficient on RCOMP in the cointegrating regression was found not to be significantly different from zero, and so were deemed not to be 'commodity-currency' countries. Accordingly, they are not listed in either the lower part of Table 2 or in Tables 3-5.

Table 2. Cointegration and Stability Tests, Real Exchange Rate and Real Commodity Prices, 1980-2002

Country	Cointegrating Parameters		R^2	Hansen Tests		
	RCOMP	*DUM*		Lc	MeanF	SupF
(1)	(2a)	(2b)	(3)	(4)	(5)	(6)
Countries Rejecting the Null Hypothesis of No Cointegration in Favor of Cointegration						
Australia	0.506 (0.122)		0.729	0.036	0.456	1.644
Bangladesh	0.327 (0.087)		0.371	0.076	0.457	1.373
Bolivia	1.164 (0.174)		0.519	0.239	2.386	8.732
Burundi	0.559 (0.088)		0.718	0.119	0.681	1.568
Ecuador	2.028 (0.339)		0.349	0.219	2.070	7.020
Iceland	0.162 (0.053)		0.409	0.123	2.314	5.197
Kenya	0.359 (0.107)		0.589	0.211	1.389	3.105
Paraguay	0.989 (0.169)		0.634	0.114	1.022	4.015
Countries Rejecting the Null Hypothesis of No Cointegration in Favor of Cointegration with a Structural Shift						
Cameroon	-0.237 (0.079)	-0.369 (0.036)	0.788			
Central African Republic	0.230 (0.058)	-0.506 (0.034)	0.909			
Côte d'Ivoire	-0.175 (0.048)	-0.338 (0.039)	0.705			
Ghana	1.270 (0.256)	-1.451 (0.260)	0.861			
Indonesia	1.169 (0.125)	-0.581 (0.086)	0.869			
Malawi	0.391 (0.135)	-0.306 (0.055)	0.699			
Mali	0.287 (0.058)	-0.494 (0.036)	0.904			
Mauritania	1.049 (0.064)	-0.257 (0.038)	0.947			
Morocco	0.709 (0.065)	0.189 (0.029)	0.854			
Niger	0.419 (0.026)	-0.460 (0.027)	0.957			
Papua New Guinea	0.366 (0.074)	-0.231 (0.037)	0.869			
Syrian Arab Republic	-0.614 (0.163)	-1.188 (0.146)	0.779			
Togo	0.297 (0.059)	-0.308 (0.030)	0.868			
Tunisia	0.164 (0.061)	-0.291 (0.024)	0.964			

Notes: The data (described in Appendices I and II) for all countries are monthly and are expressed in logarithmic form. The estimated cointegrating parameters are from the Fully Modified (FM) cointegrating regression (Phillips and Hansen (1990)): $REER = \beta_0 + \beta_1 RCOMP + \beta_2 DUM + \varepsilon$, where *REER* is the country's real effective exchange rate; *RCOMP* the national real commodity price; and *DUM is* the dummy for the shift in the cointegrating relationship; and are reported in columns 2a (for *RCOMP*) and 2b (for *DUM*); the asymptotically correct standard error of these estimates are in parentheses. All cointegrating regressions have been run using the Bartlett kernel, Andrews (1991) automatic bandwidth selector and the pre-whitened kernel estimator of Andrews and Monahan (1992). Column (3): R^2 is the regression's adjusted coefficient of determination. Columns (4-6): the 5 (10) percent critical values for the Hansen (1992) tests of parameter stability (Lc, MeanF, and SupF) are 0.623, 6.22, and 15.2 (0.497, 5.20, and 13.4), respectively. For columns 4-6, an asterisk (*) denotes statistical significance at the 5 percent level. Gregory-Hansen (1996a) tests for the presence of a regime shift in the cointegrating vector (reported in Appendix IV) reveal that the null hypothesis of no cointegration can be rejected, indicating a significant level shift in the cointegrating relation, for: the CFA franc zone countries of Cameroon, Central African Republic, Côte d'Ivoire, Mali, Niger, Senegal, and Togo (all in 1993:12); Bolivia (in 1986:01); Costa Rica (in 1998:12); Ghana (in 1983:09); Indonesia (in 1997:10); Kenya (in 1995:05); Madagascar (in 1986:04); Malawi (in 1994:08); Mauritania (in 1998:03); Mauritius (in 1986:03); Morocco (in 1992:12); Papua New Guinea (in 1995:03); Peru (in 1988:12); Syrian Arab Republic (in 1988:05); and Tunisia (in 1986:06). Accordingly, level shift dummy variables (*DUM*) have been included in the estimation of the cointegrating regressions for these countries (see lower panel above). While the null hypothesis of no cointegration could be rejected for Costa Rica, Ethiopia, Madagascar, Mauritius, Peru, Senegal, and Zambia, the coefficient on RCOMP in the cointegrating regression was found not to be significantly different from zero, and so they were deemed not to be 'commodity-currency' countries.

of Table 2.[24] All estimates of the commodity price elasticity of the real exchange rate are positive, and all cointegrating regressions have good explanatory ability, with coefficients of determination ranging between about 0.4 and 0.7.

One potential problem with time series regression models is that the estimated parameters may be unstable. In particular, the many exogenous shocks and policy changes that significantly affect small economies may cause the parameter estimates in the cointegrating relationship between each country's REER and RCOMP to change over time. Accordingly, in interpreting the relationship between these variables it is important that the long-run parameter estimates be structurally stable. To examine the hypothesis of parameter instability in the context of FM estimation of a cointegrated regression model, we again use the tests suggested by Hansen (1992). The results indicate that there is no evidence of instability in the relationship between each country's REER and the RCOMP (at the 5 percent level of significance) for any of the eight countries found to have a cointegrating relationship, as the null of parameter stability is not rejected by any of the tests (see columns (4)–(6) of the upper panel of Table 2). Accordingly, evidence of a stable cointegrating relation between the two series is found for these eight countries.

How complete is the ability of real commodity prices to explain movements in the real exchange rate of countries with commodity currencies? On average across these 22 countries, over 80 percent of the variation in the real exchange rate can be accounted for by real commodity prices (and the structural shift dummy, where appropriate), which is a surprising strong result. Clearly, movements in real commodity prices are an important driver of real exchange rates in such commodity-dependent countries.

In summary, standard cointegration tests provide only limited evidence of long-run relationships between the real exchange rate and real commodity prices. In contrast, the evidence for a cointegrating relationship between these variables, allowing for a structural shift (of unknown timing) is much more conclusive. In general, the timing of a shift in the long-run relationship between real exchange rates and real commodity prices coincides with periods of sharp revaluation of real exchange rates, arising from either nominal exchange rate revaluation and/or an associated jump in domestic and foreign inflation differentials. For two-fifths (22 of 58) of the commodity-exporting countries in our sample, the general inference to be drawn from our findings is that movements in national real exchange rates are not independent of the evolution of world real commodity prices.[25]

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[24] While the null hypothesis of no cointegration could be rejected in favor of the alternative hypothesis of cointegration for Costa Rica and Zambia, for these countries the coefficient on RCOMP in the cointegrating regression was found not to be significantly different from zero, and so were deemed not to be 'commodity-currency' countries. Accordingly, they are not listed in either the upper part of Table 2 or in Tables 3-5.

[25] For those commodity-exporting countries which could not reject the null hypothesis of no cointegration between real exchange rates and real commodity prices, it is likely that one or more highly variable factors have been omitted from the cointegrating relationship.

For those 'commodity-currency' countries that indicate that there is a long-run relationship between each country's REER and RCOMP (the 8 countries exhibiting cointegration with stable parameters and the 14 countries exhibiting cointegration with a structural shift), the value of the elasticity of each country's REER with respect to the RCOMP is of particular interest. Estimates of this elasticity range from about -0.61 (for the Syrian Arab Republic) to 2.03 (for Ecuador), with all but three of these elasticities being positive and significantly different from zero.[26] For most commodity currencies, this elasticity typically ranges between 0.2 and 0.4 (see Figure 11). The median value of the elasticity is 0.38, indicating that a 10 percent rise in real commodity prices is typically associated with a 3.8 percent appreciation of the real exchange rate of those countries with commodity currencies.

C. Commodity Currencies—Are They Real Exchange Rate Targeters?

Calvo and Reinhart (2001) point out that for commodity-exporting countries, fluctuations in real commodity prices typically require an adjustment in the real exchange rate. However, for a sample of Latin American commodity-exporting countries, Calvo, Reinhart, and Végh (1995) find that such countries may be targeting their real exchange rates, and are indexing their nominal exchange rate to the domestic price level in order to avoid losses in external competitiveness. Such a policy of real exchange rate targeting may also be implemented through domestic monetary policy, using a combination of foreign exchange reserves, open market operations and interest rate changes to dampen real exchange rate movements in the wake of real shocks. By following such 'PPP rules,' real exchange rate targeters do not accommodate real shocks by allowing their nominal exchange rate and/or relative prices to move. In this section we examine whether there is evidence that countries with commodity currencies target their real exchange rates.

For our group of 22 commodity currencies we find little to indicate that they are following such 'PPP rules,' as their real exchange rates readily move in response to real shocks. Across all countries, the mean trend decline in (the monthly rate of change) of real exchange rates is 0.2 percent per month. The volatility (measured as the standard deviation) of the rate of change of real exchange rates is much larger than the trend decline, at

[26] For three countries (Cameroon, Côte d'Ivoire, and the Syrian Arab Republic) a significant inverse relationship is found between real commodity prices and the real exchange rate, which is inconsistent with the theoretical model of Section II. In explaining this surprising result, it appears that in all three cases the estimated long-run relation has been influenced by a period of real exchange rate appreciation and declining real commodity prices (for Cameroon and Côte d'Ivoire between the mid-1980s and mid-1990s, and for the Syrian Arab Republic during most of the decade of the 1980s). In all cases the appreciating real exchange rate was driven by the use of the exchange rate as a nominal anchor; the overvaluation of the real exchange rate was subsequently removed by sharp nominal devaluations (which occurred in January 1994 (Cameroon and Côte d'Ivoire) and January 1988 (Syrian Arab Republic)). These relatively long-lived comovements appear to be being picked up as part of the long-run relation between real exchange rates and real commodity prices, rather than as part of the adjustment process (Lane and Milesi-Ferretti (2002); Reinhart and Rogoff (2002)).

Figure 11. Frequency Distribution of the Commodity Price Elasticity of the Real Exchange Rate

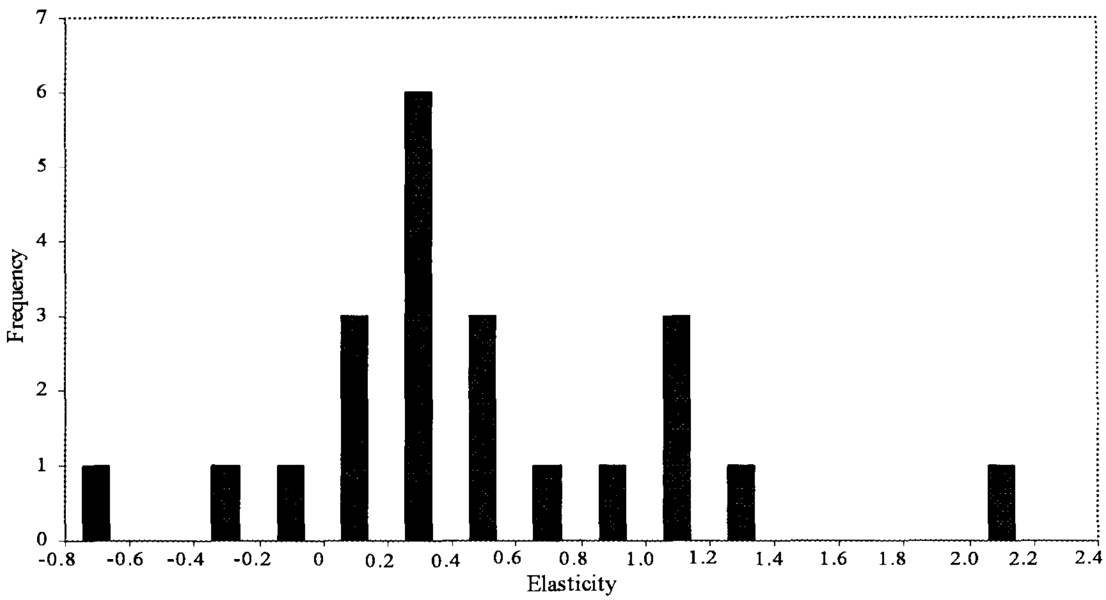
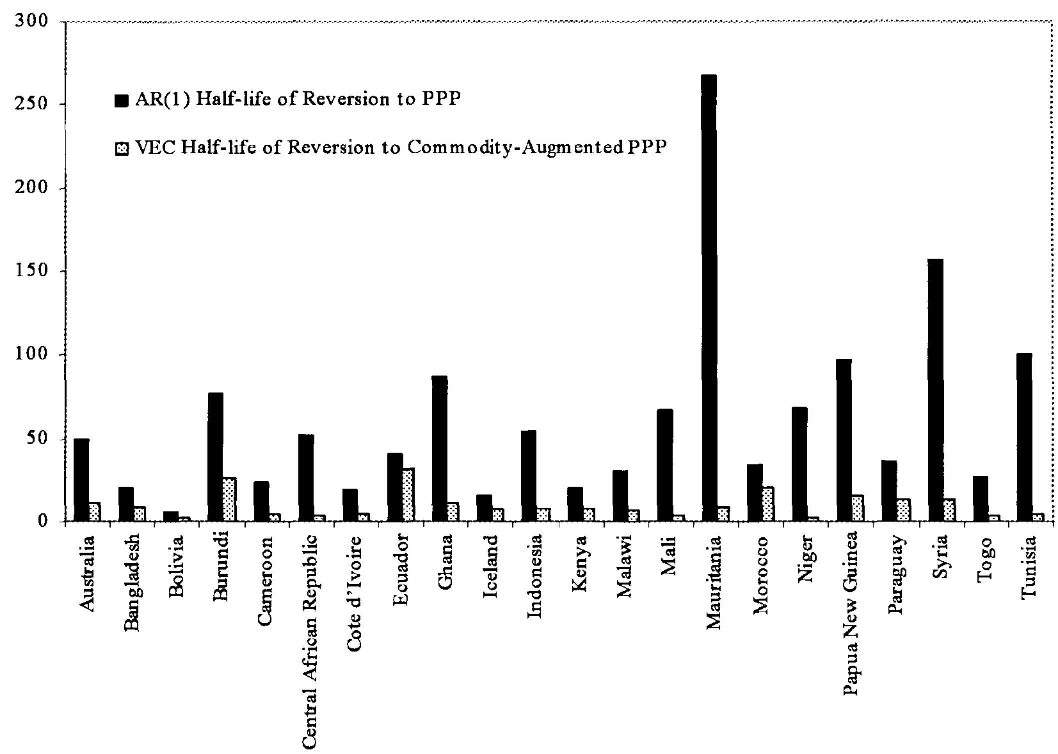


Figure 12. Half-Lives (in Months) of Reversion to PPP and Commodity-Augmented PPP, Commodity-Currency Countries



4.81 percent per month (see Table 3). Clearly, the large dispersion of real exchange rate changes dominates the relatively small secular depreciation in real exchange rates. As to real commodity prices, across all countries the mean trend decline (in the monthly rate of change) of real commodity prices is 0.28 percent per month. The volatility (measured as the standard deviation) of the rate of change of real commodity prices is much larger than the trend decline, at 3.82 percent per month. As with commodity-currency real exchange rates, the dispersion of real commodity price changes is much larger than the relatively small secular decline in real commodity prices (see also Cashin and McDermott (2002)). The decline in the real exchange rates of commodity currencies over the past 20 years tracks the fall in the real price of their commodity exports. In addition, the coefficient of variation results indicate that commodity-currency real exchange rates are, on average, about twice as variable as their real commodity prices. In summary, the information on month-to-month rates of change in the real exchange rate and real commodity prices indicates that (averaging across all commodity currencies) the real exchange rate exhibits relatively *larger* variability of monthly rates of change.

However, are these differences in the sample variances statistically significant? In answering this question we implement, *for each country found to have a commodity currency*, a nonparametric test of the equality of the variance of (the monthly rates of change) its real exchange rate and real commodity price series (Brown and Forsythe (1974) test). The results (listed in column (6) of Table 3)) indicate that the null hypothesis of equal variances is rejected, at a confidence level of at least 95 percent, for all but seven countries. Accordingly, monthly rates of change of real exchange rates are at least as variable, and in most cases more variable, than the monthly rate of change of real commodity prices. Countries with commodity currencies, be they endowed with flexible or fixed nominal exchange rate regimes, accommodate real shocks by allowing their real exchange rates to adjust. Commodity currencies are clearly not real exchange rate targeters, as they have an exchange rate policy which is directed to realigning the real exchange rate with its fundamental determinant, that is, real commodity prices.[27]

D. Exchange Rate Regimes and Commodity Currencies

When shocks to an economy are mostly real, theoretical considerations would recommend that flexible nominal exchange rates allow economies to better smooth the path of real output to such shocks, especially where domestic wages and prices are sticky. Following a negative real shock (such as a decline in the world price of a key export), a nominal depreciation raises the domestic price of exported goods to partially offset the decline in the international price, and reduces real wages in line with reduced labor demand. In contrast, countries experiencing such negative real shocks, yet with inflexible nominal

[27] See also Frankel and Saiki (2002) for a recent proposal that commodity-dependent countries should peg their currency to the world price of their dominant commodity export.

Table 3. Volatility of the Real Exchange Rate (REER) and Real Commodity Price (RCOMP), 1980-2002

Country	Std. Dev. REER	Coeff. Var. REER	Std. Dev. RCOMP	Coeff. Var. RCOMP	BF *Dispersion* REER-RCOMP (*p*-value)
(1)	(2)	(3)	(4)	(5)	(6)
Australia	0.023	23.68	0.021	9.49	0.897 (0.34)
Bangladesh	0.020	46.31	0.034	22.85	30.366 (0.00)
Bolivia	0.197	144.94	0.027	8.51	11.054 (0.00)
Burundi	0.036	27.19	0.046	9.70	14.503 (0.00)
Cameroon	0.045	40.14	0.029	10.36	3.654 (0.06)
Central African Rep.	0.046	19.11	0.039	10.74	15.42 (0.00)
Côte d'Ivoire	0.043	29.28	0.073	20.13	29.091 (0.00)
Ecuador	0.054	35.83	0.079	57.28	40.123 (0.00)
Ghana	0.114	17.00	0.037	10.51	1.824 (0.18)
Iceland	0.021	30.37	0.032	8.04	38.309 (0.00)
Indonesia	0.066	17.49	0.033	19.39	2.279 (0.13)
Kenya	0.035	78.68	0.044	15.31	10.481 (0.00)
Malawi	0.046	30.09	0.021	32.08	16.062 (0.00)
Mali	0.044	17.51	0.042	11.11	16.013 (0.00)
Mauritania	0.028	10.29	0.023	10.93	1.720 (0.19)
Morocco	0.012	10.65	0.022	9.48	47.352 (0.00)
Niger	0.051	13.83	0.063	10.65	17.145 (0.00)
Papua New Guinea	0.025	13.70	0.029	8.87	19.784 (0.00)
Paraguay	0.044	16.40	0.034	13.14	0.109 (0.74)
Syrian Arab Republic	0.047	56.37	0.069	28.02	41.925 (0.00)
Togo	0.045	22.43	0.026	9.68	1.187 (0.28)
Tunisia	0.013	8.64	0.018	11.11	54.072 (0.00)
All countries:					
Mean	0.048	32.27	0.038	15.79	
Median	0.044	23.05	0.033	10.83	
Standard deviation	0.040	30.24	0.018	11.37	

Notes: Columns (2) and (4): Std Dev is the standard deviation of the (monthly) rate of change of the series. Column (3) and (5): Coeff Var is the coefficient of variation of the (monthly) rate of change of the series. Column (6): BF is the Brown-Forsythe (1974) test statistic of the null hypothesis of equality of the variance of the (rate of change) of the REER and RCOMP series of each country; the statistic is asymptotically distributed as an F with (1, 530) degrees of freedom. The REER and RCOMP series are in logarithms and are monthly in frequency.

exchange rates, need price and wage falls to ensure that employment and output do not decline. In a seminal contribution, Mussa (1986) stressed that exchange rates behaved differently under alternative exchange rate regimes, finding that the post-Bretton Woods float of major currencies had induced large real exchange rate variability in many industrial countries. In this section we analyze the behavior of our commodity currencies, examining the importance of nominal exchange rate flexibility and relative price flexibility in driving movements in the flexible real exchange rates of commodity currencies.

Table 4 classifies our commodity currencies by exchange rate regime, using: (in column (3)) the International Monetary Fund's (1996b) de jure classification, which is based on the publicly-stated commitment of the authorities of the country in question; and (in column (2)) the de facto classification of Reinhart and Rogoff (2002), which is based on the observed behavior of market-determined real exchange rates, including that of active parallel exchange rate markets.[28] As outlined in Section IV.C, through exchange market intervention and/or monetary policy, the authorities can transform a de jure flexible exchange rate regime into a *de facto* pegged regime. Similarly, active parallel markets can transform *de jure* pegged official exchange rates into de facto flexible regimes.

Do exchange rate regimes matter for the behavior of commodity currencies and their economies? Under either the de jure or de facto classification, our commodity currencies contain a mix of currencies with flexible and fixed exchange rate regimes. We divide our commodity currencies into three groups using the Reinhart-Rogoff classification of exchange rate regimes: those which are pegs (single currency, SDR, and official basket pegs); those which display limited flexibility and/or are managed floats (including crawling peg bands and cooperative arrangements); and those which are flexible. We find that the variability of the REER of commodity currencies is similar across the various nominal exchange rate regimes (see column (4) of Table 4). This result is contrary to Mussa (1986), where industrial country real exchange rates were found to be substantially more variable during flexible nominal exchange rate regimes than they were during fixed nominal exchange rate regimes. The volatility of each country's real effective exchange rate (REER) can be decomposed into the volatility of the nominal effective exchange rate (NEER) and the volatility of the price differentials between the home country and its trading partners (RELP)—see columns (4-6) of Table 4.[29] As expected, the volatility of NEER relative to the volatility of RELP is

[28] The IMF *de jure* classification used between 1986-97 in its *Annual Report on Exchange Arrangements and Exchange Restrictions* consisted of ten categories, grouped into: pegs (1-5); limited flexibility with respect to a single currency, cooperative arrangement (6-7); more flexible arrangements, including managed floating (8-9); and independently floating (10). The Reinhart-Rogoff (2002) *de facto* classification describes exchange rate regimes as: (i) de facto pegs (including no separate legal tender and currency boards); (ii) limited flexibility (including crawling pegs and narrow crawling bands); (iii) managed floating (including wider crawling bands); (iv) freely floating; and (v) freely falling (where the annualized rate of inflation exceeds 40 percent).

[29] Volatility is measured as the standard deviation of the rate of change of each series (see Table 4).

Table 4. Volatility of Exchange Rates, Relative Prices and Real Output, 1980-2002

Country	Reinhart-Rogoff (2002) Classification of Exchange Rate Regime (de facto)	IMF Classification of Exchange Rate Regime (de jure)	Std Dev REER	Std Dev NEER	Std Dev RELP	Ratio Std Dev	Std Dev RY	BF *Dispersion* NEER-RELP (*p*-value)
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
I. Peg								
Bangladesh	Peg	Managed float	0.020	0.016	0.011	1.6	0.018	48.95 (0.00)
Cameroon	Peg	Peg	0.045	0.044	0.019	2.3	0.059	0.59 (0.44)
Central African Rep.	Peg	Peg	0.046	0.039	0.019	2.1	0.047	3.92 (0.05)
Côte d'Ivoire	Peg	Peg	0.043	0.044	0.014	3.1	0.042	0.91 (0.34)
Mali	Peg	Peg	0.044	0.042	0.027	1.6	0.045	9.41 (0.00)
Niger	Peg	Peg	0.051	0.043	0.025	1.7	0.055	8.20 (0.00)
Papua New Guinea	Peg	Peg	0.045	0.026	0.008	3.4	0.057	54.31 (0.00)
Togo	Peg	Peg	0.047	0.043	0.018	2.4	0.067	0.13 (0.72)
Mean			0.043	0.037	0.018	2.3	0.049	
Median			0.045	0.042	0.018	2.2	0.051	
Standard deviation			0.009	0.010	0.007	0.7	0.015	
II. Limited flexibility/Managed float								
Iceland	Limited flexibility	Peg	0.021	0.025	0.016	1.6	0.030	6.94 (0.01)
Morocco	Limited flexibility	Peg	0.012	0.011	0.007	1.7	0.053	1.10 (0.29)
Bolivia	Limited flexibility	Managed float	0.197	0.235	0.118	2.0	0.029	0.26 (0.61)
Indonesia	Limited flexibility	Managed float	0.066	0.067	0.013	5.1	0.048	35.04 (0.00)
Tunisia	Limited flexibility	Managed float	0.013	0.012	0.004	3.3	0.028	67.50 (0.00)
Mauritania	Limited flexibility	Managed float	0.028	0.026	0.018	1.5	0.024	5.46 (0.02)
Burundi	Managed float	Managed float	0.036	0.029	0.022	1.3	0.050	1.31 (0.25)
Kenya	Managed float	Managed float	0.035	0.034	0.015	2.3	0.020	26.60 (0.00)
Paraguay	Managed float	Independent float	0.044	0.053	0.034	1.5	0.038	15.26 (0.00)
Syrian Arab Rep.	Managed float	Peg	0.047	0.046	0.022	2.1	0.057	6.78 (0.01)
Mean			0.050	0.054	0.027	2.2	0.038	
Median			0.036	0.031	0.017	1.8	0.034	
Standard deviation			0.054	0.066	0.033	1.2	0.013	
III. Freely floating/falling								
Australia	Freely floating	Independent float	0.023	0.023	0.003	7.1	0.021	239.2 (0.00)
Ecuador	Freely falling	Managed float	0.054	0.058	0.019	3.1	0.037	29.74 (0.00)
Ghana	Freely falling	Managed float	0.114	0.122	0.026	4.7	0.036	5.26 (0.02)
Malawi	Freely floating	Managed float	0.046	0.049	0.019	2.6	0.058	11.79 (0.00)
Mean			0.059	0.063	0.017	4.4	0.038	
Median			0.050	0.054	0.019	3.9	0.037	
Standard deviation			0.039	0.042	0.010	2.0	0.015	

Notes: Col. (2): Reinhart and Rogoff's (2002) *de facto* classification of the nominal exchange rate regime. Col. (3): The IMF's *de jure* classification of the nominal exchange rate regime. In columns (2) and (3), exchange rate regimes are classified according to the regime in place for the majority of years of the 1980-2002 period. The NEER (nominal effective exchange rate, which is the trade-weighted average of bilateral exchange rates vis-à-vis trading partners' currencies); RELP (the domestic price level relative to those of trading partners, which is the differential between the domestic price level and the trade-weighted foreign price level); and REER (real effective exchange rate) indices are in logarithms (base 1995=100) and are monthly in frequency. RY (real per capita GDP growth), measured in constant 1995 units of local currency, is in logarithms and is annual in frequency. Columns (4) to (6): Std Dev is the standard deviation of the (monthly) rate of change of the series. Column (7): Ratio Std. Dev. is the ratio of the standard deviation of the (monthly) rate of change of NEER relative to the standard deviation of the (monthly) rate of change of RELP. Column (8): Std Dev is the standard deviation of the (annual) rate of change of RY. Column (9): BF is the Brown-Forsythe (1974) test statistic of the null hypothesis of equality of the variance of the (rate of change) of the NEER and RELP series of each country; the statistic is asymptotically distributed as an F with (1, 530) degrees of freedom.

smallest (largest) for the countries with pegged (flexible) exchange rate regimes, with the average ratio of volatilities almost doubling from (pegged) 2.3 to (flexible) 4.4 (see column (7) of Table 4). While the variability of the REER is similar across the various nominal exchange rate regimes, for countries with pegged nominal regimes, a larger relative share of real exchange rate variability is driven by the variability of relative prices.

Our findings of a relatively greater contribution of relative price movements in driving real exchange rate movements for commodity currencies with pegged exchange rate regimes is consistent with work by Reinhart and Rogoff (2002). They find that African CFA franc zone countries constitute the region of the world that has experienced by far the most frequent bouts of deflation in recent decades. Over the period 1970-2001, the CFA franc zone countries experience deflation (as measured by declines in the 12-month percent change in consumer prices) about 28 percent of the time. Consistent with our results in Table 4, this indicates that commodity currencies with pegged exchange rate regimes (three-quarters of which are CFA franc zone countries) are experiencing deflation (or at least inflation rates less than their trading partners) in adjusting their real exchange rates downward in response to the (typically) adverse movements in their real commodity prices.[30]

We also examine whether (for each country found to have a commodity currency) these differences in the sample variances of NEER and RELP are statistically significant, by implementing the Brown-Forsythe (1974) test of the equality of the variance of (the monthly rates of change) of NEER and RELP. The results (listed in column (9) of Table 4)) indicate that the null hypothesis of equal variances is rejected, at a confidence level of at least 95 percent, for five of the eight peg countries, seven of the ten intermediate countries, and all of the four countries with flexible exchange rate regimes. Accordingly, as expected there is evidence that the monthly rates of change of nominal effective exchange rates is more variable than the monthly rate of change of relative prices for those commodity currencies with greater flexibility in their nominal exchange rate regimes. There also appears to be no significant difference in the volatility of real output growth across the three groups of nominal exchange rate regimes (see column (8) of Table 4), as the real exchange rates of commodity currencies are sufficiently variable (either through relative prices or the nominal effective exchange rate channel) to smooth the path of output in response to real shocks.

[30] For CFA franc zone commodity currencies with pegged nominal exchange rate regimes, the share of deflationary episodes (as measured by declines in the annualized rate of change of consumer prices) for individual countries over the 1971-2001 period is: Cameroon (deflation 14 percent of the time); Central African Republic (45); Côte d'Ivoire (12); Mali (35); Niger (30); and Togo (24). In contrast, the share of deflationary episodes for the commodity currencies with flexible nominal exchange rate regimes over the same period is: Australia (zero); Ecuador (2); Ghana (3); and Malawi (1) (Reinhart and Rogoff (2002)).

E. Causality Tests

Evidence of cointegration rules out the possibility of the estimated relationship being a "spurious regression." As noted in Section IV.A, for about two-fifths of the countries in our sample, a long-run relationship between the real exchange rate and real commodity prices was found in the data. Given that cointegration has been established, then the nonstationary variable RCOMP can be thought of as encompassing the long-run component of the REER, while the residual in the cointegrating regression captures the short-run movements of the REER. It is well known that when two or more variables are cointegrated, there necessarily exists causality in *at least* one direction, and the direction of causality can be ascertained using the vector error correction (VEC) methodology suggested by Engel and Granger (1987). In the presence of cointegration, there is an error-correction representation of the relationship that implies that changes in the dependent variable are a function of the magnitude of disequilibrium in the cointegrating relationship (captured by the error-correction term), and of changes in other explanatory variables. A Wald test applied to the joint significance of the sum of the lags of each explanatory variable (testing for strict or "short-term" Granger noncausality), and a t-test of the lagged error-correction term (testing for the weak exogeneity of the variable with respect to long-run parameters, or "long-term" Granger noncausality). The bivariate VECM in REER and RCOMP will provide insight into how the long-run equilibrium is reestablished between real exchange rates and their real commodity price fundamentals.

The Granger causality tests are conducted using the VEC procedure for countries with real commodity price and real exchange rate variables that are cointegrated. In error correction form the model becomes:

$$\Delta REER_t = \eta + \sum_{i=1}^{p-1} \alpha_i \Delta REER_{t-i} + \sum_{j=1}^{p-1} \beta_j \Delta RCOMP_{t-j} + \Theta\,(REER - \kappa RCOMP)_{t-1} + e_t \qquad \textbf{(18)}$$

$$\Delta RCOMP_t = \eta' + \sum_{i=1}^{p-1} \gamma_i \Delta REER_{t-i} + \sum_{j=1}^{p-1} \delta_j \Delta RCOMP_{t-j} + \Omega\,(REER - \kappa RCOMP)_{t-1} + e'_t \qquad \textbf{(19)}$$

where: e and e′ are serially-uncorrelated disturbance terms; and the lagged error-correction term $(\text{REER} - \kappa\text{RCOMP})_{t-1}$ is the lagged residual from the cointegrating regression (of equations (15) and (16)) between REER_t and RCOMP_t, and measures the deviation from purchasing power parity in the previous period.[31] In equation (18), REER_t is Granger caused by RCOMP_t either through the lagged dynamic terms of RCOMP_t if all the β_j are not equal to

[31] The cointegrating vectors used are obtained using ordinary least squares estimation, and include a level shift dummy variable (φ_t, parameter value not reported) where the Gregory-Hansen test of Section IV.A indicated equation (16) was the appropriate cointegrating regression.

zero, or through the lagged error-correction term if Θ is nonzero. Similarly, $RCOMP_t$ is Granger caused by $REER_t$ either through the lagged dynamic terms of $REER_t$ if all the γ_i are not equal to zero, or through the lagged error-correction term if Ω is nonzero (equation (19)). The speeds of adjustment parameters (Θ and Ω) indicate how quickly the system returns to its long-run equilibrium after a temporary departure from it.

For those countries with commodity currencies, the results of the causality analysis using the VEC procedure are set out in Table 5. The *t*-test results support the hypothesis that errors from the cointegrating relationship significantly influence changes in the real exchange rate, but there is much weaker evidence that they affect the real price of commodities. For six countries (such as Australia and Togo), the Wald tests indicate there is evidence (at the 10 percent level of significance) that short-run movements in RCOMP help predict (Granger cause) part of the short-run movement in REER. For *all* 22 countries, with Θ less than zero (which ensures error correction) and statistically significant, a positive (negative) disequilibrium term $(REER - \kappa RCOMP)_{t-1}$ will ensure that REER declines (rises) toward its long-run equilibrium path.[32] These results imply that RCOMP was the initial receptor of exogenous shocks to the long-term relationship, and REER had to adjust to reestablish the long-run equilibrium. Accordingly, we find that real exchange rates adjust to eliminate errors in the cointegrating regression, and there is evidence in support of the notion of rising real commodity prices leading to increasing (appreciating) real exchange rates. In line with our expectations, these results indicate that in the Granger-causality sense, RCOMP (as the more exogenous variable) predominantly leads (rather than lags) REER, and that the latter undertakes the short-run adjustment necessary to reestablish the long-run equilibrium.[33]

F. The PPP Puzzle and Commodity Currencies

Although the central issue discussed in this paper is the role played by real commodity prices in driving movements in the real exchange rate, our econometric results also appear to offer a potential resolution of the well-known "purchasing power parity (PPP) puzzle" (Rogoff (1996)). This puzzle concerns the finding of many researchers that the speed of mean reversion of real exchange rates is too slow to be consistent with PPP, which is the

[32] This finding of the coefficient on the error-correction term being appropriately negative and significantly different from zero also means that econometric specifications based on first differences of the variables alone will probably be ignoring useful information about the parity-reverting properties of the real exchange rate.

[33] In Table 5 there is much less evidence of rising REER leading to increasing RCOMP, either through the lagged dynamic terms of REER (for only 4 of 22 countries is there evidence that short-run movements in REER Granger-cause short-run movements in RCOMP) or the lagged error-correction term (as for half the countries Ω is typically negative or not significantly different from zero, which does not ensure error correction). For 12 countries there is evidence of bidirectional causality through the error-correction term, although the parameter on the error-correction term in equation (19) is in these cases relatively small, indicating that the real exchange rate bears most of the burden of adjustment to long-run equilibrium.

Table 5. Causality Between Real Exchange Rate (REER) and Real Commodity Price (RCOMP) Using VEC Approach

			H_0: RCOMP does not Granger-cause REER		H_0: REER does not Granger-cause RCOMP	
Country	Half-life of Shock to Commodity-Augmented PPP (months)	Lag Order	β_j = 0: χ^2-stat (*p*-value)	Θ (*t*-stat)	γ_i = 0: χ^2-stat (*p*-value)	Ω (*t*-stat)
(1)	(2)	(3)	(4)	(5)	(6)	(7)
Australia	11.40	2	5.189 (0.07)	-0.059 (-3.12)	0.853 (0.65)	-0.001 (-0.05)
Bangladesh	8.65	1	1.323 (0.25)	-0.077 (-4.09)	0.316 (0.57)	-0.014 (-0.41)
Bolivia	2.38	1	0.693 (0.41)	-0.253 (-5.96)	0.886 (0.35)	-0.002 (-0.41)
Burundi	26.31	1	1.074 (0.30)	-0.026 (-1.90)	0.001 (0.98)	0.034 (3.26)
Cameroon	3.95	1	1.357 (0.24)	-0.161 (-4.82)	0.086 (0.77)	-0.038 (-1.77)
Central African Rep.	3.56	1	2.819 (0.09)	-0.177 (-5.42)	4.984 (0.03)	0.043 (1.76)
Côte d'Ivoire	4.09	1	2.019 (0.16)	-0.156 (-5.26)	0.030 (0.86)	-0.106 (-3.72)
Ecuador	31.16	1	1.238 (0.27)	-0.022 (-1.78)	5.693 (0.02)	0.066 (3.80)
Ghana	11.01	1	0.390 (0.53)	-0.061 (-3.18)	2.290 (0.13)	0.019 (3.13)
Iceland	8.10	1	0.705 (0.40)	-0.082 (-3.55)	0.214 (0.64)	0.035 (0.99)
Indonesia	7.80	1	0.166 (0.68)	-0.085 (-3.42)	0.278 (0.59)	0.035 (3.07)
Kenya	7.44	1	4.549 (0.03)	-0.089 (-3.86)	0.580(0.45)	-0.022 (-0.71)
Malawi	6.94	1	2.954 (0.09)	-0.095 (-4.14)	1.284 (0.26)	0.033 (3.37)
Mali	3.07	1	7.307 (0.01)	-0.202 (-6.95)	10.298 (0.00)	0.031 (1.21)
Mauritania	8.54	1	0.001 (0.97)	-0.078 (-3.26)	0.138 (0.71)	0.056 (2.81)
Morocco	20.66	2	3.878 (0.14)	-0.033 (-1.85)	1.976 (0.37)	0.123 (3.94)
Niger	2.43	1	1.618 (0.20)	-0.248 (-6.33)	0.134 (0.71)	0.121 (2.34)
Papua New Guinea	15.77	2	1.888 (0.39)	-0.043 (-2.02)	0.409 (0.82)	0.048 (2.15)
Paraguay	14.09	1	0.025 (0.87)	-0.048 (-2.62)	0.051 (0.82)	0.037 (2.75)
Syrian Arab Republic	13.51	1	0.327 (0.57)	-0.050 (-3.68)	0.156 (0.69)	-0.043 (-2.13)
Togo	2.99	1	5.888 (0.02)	-0.207 (-5.73)	8.204 (0.00)	0.042 (2.03)
Tunisia	4.03	2	4.232 (0.12)	-0.158 (-6.43)	2.890 (0.24)	0.038 (1.02)
All countries:						
Mean	5.97			-0.110		0.024
Median	7.95			-0.084		0.051
Standard deviation				0.073		0.035

Notes: See equations 18 and 19. The cointegrating vectors used are obtained using ordinary least squares estimation, and include a level shift dummy variable (parameter value not reported) where the Gregory-Hansen test of Section IV.A indicated was appropriate. The lag length of column (3) is determined by minimizing the Akaike Information criterion. The implied half-life of the shock to commodity-price-augmented PPP (column (2)) is calculated as follows. The time (T) required to dissipate x percent (in this case, 50 percent) of a shock is determined according to: $(1-\Theta)^T = (1-x)$, where Θ is the coefficient of the error-correction term and T is the required number of periods (months). Multivariate Lagrange Multiplier (LM) tests for serial correlation (with the order of serial correlation tested being one more than the optimal lag length of the VEC model) indicate that there is little evidence of residual serial correlation.

proposition that exchange rates are determined by movements in relative prices.[34] In summarizing the results from studies using long-horizon data, Froot and Rogoff (1995) and Rogoff (1996) report the current consensus in the literature that the half-life of a shock (the time it takes for the shock to dissipate by 50 percent) to the real exchange rate is about three to five years, implying a slow speed of reversion to (constant) parity of between 13 to 20 percent per year. Such a slow speed of reversion to purchasing power parity is difficult to reconcile with nominal rigidities (where one would expect substantial parity-reversion over one to two years), and is also difficult to reconcile with the observed large short-term volatility of real exchange rates.

A potential solution to Rogoff's (1996) PPP puzzle may lie in identifying a (real) shock that is both sufficiently volatile and persistent to rehabilitate the purchasing power parity approach to real exchange rate determination (Chen and Rogoff (2002)). Previous work indicates that fluctuations in world commodity prices would certainly fit the bill as being a source of real shocks that are both highly persistent and rather volatile (Cashin, McDermott, and Liang (2000); Cashin, McDermott, and Scott (2002); Cashin and McDermott (2002)). Accordingly, in this section we will examine whether real commodity prices are an important variable in accounting for medium- to long-term deviations of 'commodity-currency' real exchange rates from purchasing power parity. We do so after controlling for real shocks, by incorporating real commodity prices as a determinant of the equilibrium real exchange rate of commodity currencies, and then examine the persistence of shocks to real exchange rates in reverting to their commodity-price-augmented equilibria.

To examine the extent of persistence in 'commodity-currency' real exchange rates, we begin by estimating a standard first-order autoregressive model (or Dickey-Fuller regression), without controlling for commodity prices, and focus on the magnitude of the least squares estimates of the autoregressive parameter.[35] Across all countries, the median half-life of parity reversion is 36 months for our sample of 58 commodity-dependent countries, while for the 22 'commodity currencies' the median half-life of parity reversion is

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[34] PPP implies that at equilibrium exchange rates, foreign currencies should possess the same purchasing power. If PPP is true, then in response to random monetary or real shocks, movements in the real exchange rate should be stationary and induce only temporary deviations from PPP. Commodity arbitrage would act as an error-correction mechanism to force the local currency price of a bundle of domestic goods into line with the foreign currency price of a common bundle of foreign goods.

[35] It should also be noted that as the data on prices used to construct the IMF's real effective exchange rate series are measured as indices relative to some base period (rather than price levels), what is being examined here is the relative version of PPP (in levels) and not the absolute version of PPP. If relative PPP holds in the long run, then the real exchange rate will revert to its (constant) average level.

slightly longer at 44 months.[36] These results are consistent with Rogoff's (1996) consensus of half-lives of parity reversion of between 36 to 60 months (three to five years).[37]

Next we turn to the results from our VEC model, which provides information on the speed with which real exchange rates adjust to re-establish their long-run equilibrium relationship with real commodity prices (see column 5 of Table 5). The magnitude of Θ (the coefficient on the error-correction term in equation (18)) indicates that for some countries (such as Australia and Papua New Guinea) only about 5 percent of the deviation of the REER from long-run equilibrium is eliminated in one month (implying a half-life of parity deviation of about 13½ months), while for other countries (such as the Central African Republic and Tunisia) about 16 percent of the deviation is eliminated in one month (implying a half-life of parity deviation of about 4 months), a very rapid speed of adjustment.[38, 39] For each of the 22 'commodity-currency' countries, the half-life of the reversion of the real exchange rate to its (constant) long-run *average* level and to its commodity-dependent long-run *equilibrium*

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[36] The half-life is the length of time it takes for a unit impulse to dissipate by half. The least squares estimate of the half-life of a shock (HLS) is calculated using the formula: HLS = ABS(log(1/2)/log(α)), where α is the autoregressive parameter derived from the least squares regression. These half-life results are comparable to those obtained by Cheung and Lai (2000a) using least squares estimation on monthly bilateral (post-Bretton Woods) dollar real exchange rates for developed countries, which calculated an average half-life of 3.3 years. In addition, supporting the earlier work of Edwards and Savastano (1999), Cheung and Lai (2000b) report that a finding of stationarity in real exchange rates is more likely in developing countries (particularly Latin American countries) than for developed countries, and that most of the half-lives of parity deviation for developing countries are less than three years.

[37] However, for most commodity-currency countries the associated 90 percent confidence intervals for the half-life of PPP deviations are quite wide and encompass half-lives that are inconsistent with PPP holding in the long run (given that the confidence interval includes infinity as the upper bound). This indicates that there is a high level of uncertainty about the "true" value of the half-life of PPP deviations—while the median half-life may be about three years in duration, these confidence intervals that are typically so wide that the point estimates of half-lives from the AR(1) regressions provide virtually no information regarding the true duration of the half-lives of parity reversion.

[38] The time (T) required to dissipate x percent of a shock is determined according to: $(1-\Theta)^T = (1-x)$, where Θ is the coefficient of the error-correction term and T is the required number of periods (months).

[39] While the estimated long-run effect of RCOMP on the REER can be quite large, the short-run impact is typically weaker. For example, in the case of Ecuador a 10 percent increase in RCOMP would raise the REER in the long run by 20.3 percent, but in the short run by only 0.022*(20.3) = 0.45 percent. In contrast, in the case of Iceland a 10 percent increase in RCOMP would raise the REER in the long-run by 1.62 percent, but in the short-run by only 0.082*(1.62) = 0.13 percent.

level are set out in Figure 12. The speed of reversion to the commodity-dependent equilibrium real exchange rate is, for all countries, much faster than the speed of reversion to the average real exchange rate.[40]

Averaging across all 'commodity-currency' countries, the median error correction on real exchange rates is about 8 percent per month; on commodity prices it is about 5 percent per month. The elimination of 8 percent of the deviation of the real exchange rate from its equilibrium level per month is the equivalent of a median half-life of parity deviation of about 8 months, which is much faster than the typical half-life (of about three to five years) reported in the simple PPP-based regressions analyzed above (Rogoff (1996)). That is, while the real exchange rate of those countries with commodity currencies has a slow reversion to its *average* level (the median half-life of parity deviations is 44 months), it has a much faster speed of adjustment towards its *equilibrium* level (the median half-life of parity deviations is about 8 months), where that equilibrium depends on the evolution of real commodity prices as a fundamental determinant of the real exchange rate. These results indicate that, particularly for commodity-dependent developing countries, controlling for the influence of real commodity prices on the real exchange rate is an important channel by which to reduce the measured persistence of real exchange rate shocks.

V. Conclusions

In this paper we examined the evidence for a real commodity price explanation of movements in the real exchange rates of 58 commodity-dependent countries over the period 1980-2002. For about two-fifths of the commodity-exporting countries we find robust evidence in support of the long-run comovement of national real exchange rate and real commodity price series. The long-run real exchange rate of this group of 'commodity currencies' are time-varying, being dependent on movements in real commodity prices. For all commodity currencies, there is little evidence that they target the level of their real exchange rates, as the relative volatility of national real exchange rates exceeds that of national real commodity prices. In addition, weak exogeneity tests carried out within a vector error correction framework indicate highly significant causality running from real commodity prices to the real exchange rate—it is typically the real exchange rate that adjusts to restore the long-run equilibrium relationship with real commodity prices. The commodity currencies are found to exhibit extremely rapid half-lives of adjustment of real exchange rates to

[40] In comparison with the relatively slow adjustment speed of real exchange rates to parity typically found for developed countries, nominal rigidities appear to be less important for countries with commodity currencies (which are predominantly developing countries). This relatively fast adjustment of wages and nontraded goods prices for commodity currencies is consistent with the relatively small formal sector of developing countries in comparison with that of developed countries, and with developed countries' relatively larger share of nontraded goods prices in domestic prices (see Baffes, Elbadawi, and O'Connell (1999)).

commodity-price-augmented purchasing power parity of about eight months. Such a short half-life of mean reversion to commodity-price-augmented PPP is certainly sufficiently rapid to resuscitate purchasing power parity as a key determinant of real exchange rates in commodity-dependent countries. These estimates also cast doubt on the universality of Rogoff's (1996) consensus estimate of the half-life of the reversion of real exchange rates to purchasing power parity of about three to five years. As presciently conjectured by Keynes (1930), for countries with 'commodity currencies,' controlling for a major source of real shocks (movements in real commodity-export prices) provides much stronger empirical evidence in support of purchasing power parity-based approaches to exchange rate determination. For commodity currencies, movements in real commodity prices are an important (and hitherto missing) piece of the PPP puzzle.

The results in this paper have an important implication for monetary and exchange rate policies in countries with commodity currencies. One of the stylized facts in the international finance literature is that the behavior of the real exchange rate is dependent on nominal exchange rate regimes (Mussa (1986)). Our results indicate that this is not so, as among our group of commodity currencies are countries with both fixed and flexible nominal exchange rate regimes. Accordingly, we argue that it is the nature of real shocks to the economy which determine the behavior of real exchange rates, rather than the type of nominal exchange rate regime. Where real shocks are the dominant influence on real exchange rates, commodity-dependent countries need to have either flexible nominal exchange rate regimes (which facilitate the slow change of relative inflation rates) or flexible wages and prices (which facilitate the maintenance of nominal exchange rate pegs).

Description of the Data

The data are of monthly frequency, for the period 1980:01-2002:03. The 58 potential commodity-currency countries in our sample are listed in Appendix III. The primary data sources are the IMF's *International Financial Statistics* (IFS) and *Information Notice System* (INS). Below we provide below a description of the series.

REER: Trade-weighted measure of the seasonally adjusted, CPI-based real effective exchange rate (base 1990=100); obtained from the IMF's INS.

NCOMP: The nominal commodity export price index for each country (base 1990=100, seasonally adjusted) has been calculated using UN COMTRADE data on the (1990-99 average) share of each commodity in total primary commodity exports, and the IMF's (U.S. dollar-based) data on world commodity prices (taken from the IMF's IFS). The derivation of this index is described in detail in Appendix II.

RCOMP: The real commodity export price index is calculated by: deflating each country's NCOMP by the IMF's index of the unit value of developed country manufactured exports (MUV).

MUV: Unit value index (in U.S. dollars) of manufactures exported by 20 developed countries, with country weights based on the countries' total 1995 exports of manufactures (base 1995=100); obtained from the IMF's IFS.

Construction of the Country-Specific Nominal Price Indices of Commodity Exports

The country-specific nominal export price indices (NCOMP) for the period 1980:01-2002:03 were constructed as set out below.

For each country, we calculate the 1991-99 average total value of primary commodity exports; the 44 individual nonfuel commodity weights are calculated by dividing the 1991–99 average value of each individual commodity export by the 1991-99 average total value of primary commodity exports. All commodity weights are gross export weights as found in the World Bank's World Integrated Trade Solution (*WITS*), which supplies UN *COMTRADE* data provided by the UN Statistical Department. Once the country-specific commodity export weights are established, these weights are held fixed over time and are used to weight the individual (U.S. dollar-based) price indices of the same commodities—taken from the IMF's IFS—to form, for each country, a geometric weighted-average index of (U.S. dollar-based) nominal commodity export prices (base 1990=100). The national index of nominal commodity export prices are then seasonally adjusted using the X11.2 variant of the Census Method 11 procedure.

Nominal Commodity Prices

The 44 nonfuel commodities used in the calculation of the national commodity price indices are: aluminum, bananas, beef, coal, cocoa, coconut oil, coffee, copper, cotton, fish, fish meal, gold, groundnut oil, groundnuts, hardwood logs, hides, iron, lamb, lead, maize, natural rubber, nickel, palm oil, palm kernel oil, phosphate rock, platinum, potash, rice, shrimp, silver, softwood logs, softwood sawn, soy meal, soy oil, soybeans, three types of sugar, sun/safflower oil, tea, tin, tobacco, wheat, wool, uranium, and zinc.

Below is a complete list of the 44 nonfuel commodity prices used in the construction of the geometric weighted index of commodity export prices, their sources, and a brief description. The individual commodity prices selected are the closest to a world price included in the IFS, and were indexed themselves before being weighted by the 1991-1999 average share of total primary commodity exports. The individual nonfuel commodities are selected partly due to their importance and partly due to data availability, and are:

- **Aluminum**, London Metal Exchange accessed through Bloomberg, standard grade, spot price, minimum purity 99.5 percent, c.i.f. U.K.

- **Bananas**, from January 1997 Sopisco News; prior to that USDA. The Sopisco price is Central America and Ecuador first class quality tropical pack, average of Chiquita, Dole and Del Monte, U.S. importer's price f.o.r. U.S. ports. The previous USDA price was the market price at the Philadelphia market.

- **Beef**, Urner Barry's The Yellow Sheet, Australian and New Zealand frozen boneless, 85 percent visible lean cow meat, U.S. import price f.o.b. port of entry.

- **Coal**, World Bank, Australian thermal, 12,000 btu/lb, less than 1.0 percent sulfur, 14 percent ash, f.o.b. piers, Newcastle/Port Kembla. Prior to 1982 is the percentage change in the unit value of Australian coal and briquettes exports, taken from *IFS*.

- **Cocoa**, Financial Times, International Cocoa Organization Daily price which is an average of the three nearest active futures trading months in the New York Cocoa Exchange at noon and the London Terminal market at closing, c.i.f. U.S. and European ports.

- **Coconut Oil**, Oil World Weekly ISTA Mielke GmbH, Philippine/Indonesian bulk c.i.f. Rotterdam.

- **Coffee**, International Coffee Organization (New York) price; is a weighted average of Arabica (other milds), and Robusta coffee prices, equally weighted.

- **Copper**, London Metal Exchange accessed through Bloomberg, grade A cathodes, spot price, c.i.f. European ports.

- **Cotton**, Cotton Outlook, Middling 1-3/32 inch staple, Cotlook 'A' Index, average of the cheapest five of sixteen styles, c.i.f. North Europe.

- **Fish**, Norstat. Prices after 1987 are for farm fresh salmon—prior to 1988 the price is an export unit value index of all Norwegian fish exports.

- **Fish Meal**, Oil World Weekly ISTA Mielke GmbH, any origin, 64-65 percent protein, c.i.f. Hamburg.

- **Gold**, Bloomberg, London Bullion Market Association PM fixed price.

- **Groundnut Oil**, Oil World Weekly ISTA Mielke GmbH, any origin, c.i.f. Rotterdam.

- **Groundnuts**, Oil World Weekly ISTA Mielke GmbH, U.S. Runners, c.i.f. Rotterdam.

- **Hardwood Logs**, World Bank, Malaysian, meranti, Sarawak best quality, sale price charged by importers, Japan.

- **Hides**, Wall Street Journal Previous to November 1985 US Bureau of Labor Statistics, U.S. price, Chicago packer's heavy native steers, over 53 lbs., wholesale dealer's price (formerly over 58 lbs.), f.o.b. shipping point.

- **Iron Ore**, Companhia Vale do Rio Doce, Brazilian Carajas fines, standard sinterfeed, 62.65 percent iron, contract price to Europe, f.o.b. Ponta da Madeira.

- **Lamb**, National Business Review, New Zealand medium fat content, U.K. prices.

- **Lead**, London Metal Exchange accessed through Bloomberg, 99.97 percent pure, spot price, c.i.f. European ports.

- **Maize**, USDA Grain and Feed Market News, U.S. No.2 yellow, prompt shipment, f.o.b. Gulf of Mexico ports.

- **Natural Rubber**, Financial Times, Malaysian No.1 RSS, prompt shipment, f.o.b. Malaysian/Singapore ports.

- **Nickel**, London Metal Exchange accessed through Bloomberg, melting grade, spot price, c.i.f. Northern European ports.

- **Palm Oil**, Oil World Weekly ISTA Mielke GmbH, Malaysian/Indonesian, c.i.f. Northwest European ports.

- **Palm Kernel Oil**, Malaysian, c.i.f. Rotterdam.

- **Phosphate Rock**, World Bank, Moroccan, 70 percent BPL, contract, f.a.s. Casablanca.

- **Platinum**, Thomson Financial DataStream, London Free Market price.

- **Potash**, Fertilizer Markets, standard grade of potassium chloride, f.o.b. Vancouver.

- **Rice**, USDA Rice Market News, Thai white milled, 5 percent broken, f.o.b. Bangkok.

- **Shrimp**, U.S. frozen, 26/30 count, wholesale New York.

- **Silver**, London Bullion Market Association.

- **Softwood Logs**, USDA Forest Service Pacific Northwest Research Station, average export price of Douglas-fir, Western Hemlock and other softwoods exported from Washington, Oregon, northern California and Alaska. Prior to 1982 is monthly percentage change in softwood, Western Canada, lumber prices, for 2x4 random length.

- **Softwood Sawn**, USDA Forest Service Pacific Northwest Research Station, average export price of Douglas-fir, Western Hemlock and other sawn softwoods exported from Washington, Oregon, northern California and Alaska. Prior to 1982 is monthly percentage change in softwood, Western Canada, lumber prices, for 2x4 random length.

- **Soy Meal**, Oil World Weekly ISTA Mielke GmbH, Argentine 45/46 percent protein, c.i.f. Rotterdam.

- **Soy Oil**, Oil World Weekly ISTA Mielke GmbH, Dutch c.i.f. ex-mill.

- **Soybeans**, Oil World Weekly ISTA Mielke GmbHm, U.S. c.i.f. Rotterdam.

- **'Free market' Sugar**, International Sugar Organization accessed through Bloomberg, International Sugar Organization price, average of the New York CSCE contract No. 11 spot price and the London daily price, f.o.b. Caribbean ports.

- **U.S. Sugar**, Wall Street Journal, CSCE No.14 contract, nearest futures position, c.i.f. New York.

- **E.U. Sugar**, E.U. Office in Washington D.C., E.U. import price, unpacked sugar, c.i.f. European ports. Negotiated price for sugar from ACP countries to E.U. under the Sugar Protocol.

- **Sun/Safflower Oil**, Oil World Weekly ISTA Mielke GmbHm, any origin, ex-tank Rotterdam.

- **Tea**, Reuters, Mombasa auction price for best PF1, Kenyan tea; prior to July, 1998 the London auction price.

- **Tin**, London Metal Exchange accessed through Bloomberg, standard grade, spot price, c.i.f. European ports.

- **Tobacco**, USDA (Foreign Agricultural Service), U.S. import unit value of general unmanufactured tobacco.

- **Uranium**, NUEXCO Exchange Value, Restricted price for U_3O_8, US$ per pound.

- **Wheat**, USDA Grain and Feed Market News, U.S. No.1 hard red winter, ordinary protein, prompt shipment, f.o.b. Gulf of Mexico ports.

- **Wool**, Australian Wool Exchange, weighted average of fine (19 micron) and coarse (23 micron) wool prices, with weights being their historic trading volumes.

- **Zinc**, London Metal Exchange accessed through Bloomberg, high grade, 98 percent pure, spot price, c.i.f. European ports.

The lack of a sufficiently long time series of data for several important individual commodities precluded them from being included in the final construction of the national commodity export price indices. As a result, while the national commodity export price indices will be highly correlated with the true national commodity export price indices, the omissions will tend to bias toward zero the relationship between the commodity export price index and the real exchange rate. In particular, inadequate price series for the following commodities preclude them from being included in the final construction of the export-price indices: barley, cashews, cobalt, diamonds, hardwood sawn, olive oil, poultry, and swine meat.

- For the analysis of the exports of Norway, Indonesia, Mexico (the three largest non-OPEC oil producers) and the Syrian Arab Republic, the following fuel commodities were added to those above to form the national commodity export price indices:

- **Petroleum**, spot crude, average of U.K. Brent (light), Dubai (medium), and West Texas Intermediate (heavy), equally weighted.

- **Natural Gas**, Russian border price in Germany (World Gas Intelligence, New York). Prior to 1985 is percentage change in the U.S. Department of Energy's purchase price (Bloomberg).

Robustness Checks of Nominal Commodity Export Price Series

Finally, for Australia, Canada, and New Zealand the commodity export price indices constructed here can be compared to those calculated by central banks and national commercial banks. The (U.S. dollar-based) Australian nominal commodity export-price index (for all commodities) is available from 1982:07-2002:3 (calculated by the Reserve Bank of Australia); the (U.S. dollar-based) Canadian nominal commodity export price index (for all nonenergy commodities) is available for the full sample period 1980:01-2002:3 (calculated by the Bank of Canada); and the (U.S. dollar-based) New Zealand commodity export price index (for all commodities) is available from 1986:01- 2002:03 (calculated by the Australia and New Zealand Bank). The constructed and official nominal commodity export price indices are very similar for Australia and Canada, while the constructed and bank indices for New Zealand differ somewhat, due to the exclusion of dairy products from the constructed index. These results provide some comfort as to the accuracy of the constructed nominal commodity export price indices for the other countries in our sample.

Potential Commodity-Dependent Countries

The IMF's *World Economic Outlook* classifies countries into groups, based on certain criteria (see IMF (1996a)). The groups are (for non-developing countries): *industrial countries*; and *countries in transition*. Developing countries are classified by their predominant export as: *primary product exporters* (those countries whose exports of agricultural and mineral primary products (Standard Industrial Trade Classification 0, 1, 2, 4, 68) accounted for at least 50 percent of their total export earnings in 1988-92); *fuel exporters* (those countries whose exports of fuel products (SITC 3) accounted for at least 50 percent of their total export earnings in 1988-92); *exporters of manufactures* (those countries whose exports of manufactures (SITC 5 to 8, excluding 68) accounted for at least 50 percent of their total export earnings in 1988-92); *exporters of services and recipients of factor income and private transfers* (those countries whose average income from services, factor income, and worker's remittances accounted for at least 50 percent of their total export earnings in 1988-92); and countries with a *diversified export base* (those developing countries whose export earnings were not dominated by any of the categories mentioned above).

Countries

The 58 countries included in our sample are listed below using the IMF's *World Economic Outlook* classification. Included for each country (derived from UN *COMTRADE* data) are: the years available of export data that could be averaged; and total exports of primary commodities as a percent of total exports of all goods. For example (see below), Argentina had annual UN COMTRADE data for the period 1991-99 which was used to derive the period-average share of individual commodity exports in total commodity exports; on average over the 1991-99 period, commodity exports comprised 41 percent of its total exports. The 58 potential commodity-currency countries are listed below, by geographic region, in Table 6.

Primary-product-exporting developing countries—Argentina (1991-99, 41); Bolivia (1991-99, 56); Burundi (1993-99, 97), Chile (1991-99, 58); Côte d'Ivoire (1995-99, 65), Ethiopia (1993-99, 71); Ghana (1992, 1996-99, 72), Guatemala (1991-99, 49); Honduras (1991-99, 67); Madagascar (1991-99, 39); Malawi (1991, 1994-95, 90), Mali (1995-99, 85); Mauritania (1995-98, 64); Myanmar (1991-92, 52); Nicaragua (1991-99, 69); Niger (1995-99, 67); Papua New Guinea (1991-93, 1995-99, 59); Paraguay (1991-99, 79); Peru (1991-99, 69); St. Vincent and the Grenadines (1993-99, 72); Sudan (1992-99, 44); Suriname (1991-92, 1994-98, 86); Tanzania (1997-99, 59) ; Togo (1991, 1995-99, 84); Uganda (1994-99, 84); Zambia (1993-97, 88); and Zimbabwe (1991-99, 54).[41]

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[41] Of the 43 primary-product-exporting developing countries listed by the *World Economic Outlook* (IMF 1996)), eight could not be included in our sample due to the absence of data on their real effective exchange rate (Afghanistan, Equatorial Guinea, Guinea, Guinea-Bissau, Liberia, São Tomé and Príncipe, Somalia, and Vietnam) and 8 could not be to be included due to the absence of relevant commodity-price data (Botswana, Chad, the Democratic Republic of Congo, Guyana, Namibia, Rwanda, Solomon Islands, and Swaziland).

Diversified export base developing countries—Bangladesh (1991-99, 8); Brazil (1991-99, 35); Cameroon (1994-99, 53); Central African Republic (1993-98, 43); Colombia (1991-99, 40); Costa Rica (1991-99, 31); Dominica (1991, 1993-99, 31); Ecuador (1991-99, 49); India (1991-99, 31); Indonesia (1991-99, 43); Kenya (1991-99, 45); Malaysia (1991-99, 13); Mauritius (1991-99, 27); Mexico (1991-99, 15); Morocco (1991-99, 14); Mozambique (1994-99, 26); Pakistan (1991-99, 12); Philippines (1991-99, 10); Senegal (1991-99, 26); South Africa (1991-99, 38)[42]; Sri Lanka (1991-99, 20); Syrian Arab Republic (1992, 1994-99, 74); Thailand (1991-99, 16); Tunisia (1991-99, 8); Turkey (1991-99, 8); and Uruguay (1991-99, 32).[43]

Commodity-exporting industrial countries—Australia (1991-99; 54); Canada (1991-99, 15); Iceland (1991-99, 56); New Zealand (1991-99, 35); and Norway (1991-99, 63).

[42] Due to underreporting of gold exports in the 1991-99 sample period, the commodity price index for South Africa has been amended so that gold exports reflect their historical 2:1 ratio to diamond exports.

[43] Of the 30 diversified export base developing countries listed by the *World Economic Outlook* (IMF 1996)), 4 could not be used due to the absence of data on their real effective exchange rate (Comoros, the Lao People's Democratic Republic, Netherlands Antilles, and Sierra Leone).

Table 6. Potential Commodity-Currency Countries, by Region

Sub-Saharan Africa	Asia-Pacific	Middle East and North Africa	Western Hemisphere	Europe
Burundi	***Australia***	*Morocco*	Argentina	***Iceland***
Cameroon	*Bangladesh*	*Syrian Arab Republic*	*Bolivia*	**Norway**
Central African Republic	India	*Tunisia*	Brazil	
Côte d'Ivoire	*Indonesia*	Turkey	**Canada**	
Ethiopia	Malaysia		Chile	
Ghana	Myanmar		Colombia	
Kenya	**New Zealand**		Costa Rica	
Madagascar	Pakistan		Dominica	
Malawi	*Papua New Guinea*		*Ecuador*	
Mali	Philippines		Guatemala	
Mauritania	Sri Lanka		Honduras	
Mauritius	Thailand		Mexico	
Mozambique			Nicaragua	
Niger			*Paraguay*	
Senegal			Peru	
South Africa			St. Vincent & Grenadines	
Sudan			Suriname	
Tanzania			Uruguay	
Togo				
Uganda				
Zambia				
Zimbabwe				

Notes: **Industrial** countries are denoted in bold. Those countries found to have *commodity currencies* (those which exhibit a long-run relationship between their real effective exchange rate and their real commodity export prices, and where the coefficient on real commodity-export prices in the cointegrating regression was found to be significantly different from zero) are denoted in italics.

Cointegration Tests: Real Exchange Rate and Real Commodity Prices, Commodity-Exporting Countries, 1980-2002

Country (1)	$Z(t)$ (2)	$Z(\alpha)$ (3)	$Z(t)$* (4)	Shift date (5)
Argentina	-2.05	-8.82	-3.84	
Australia	-3.63*	-24.76*	-3.53	
Bangladesh	-3.45*	-23.17*	-3.84	
Bolivia	-6.07*	-64.24*	-6.21*	[1986:01]
Brazil	-2.61	-13.86	-3.29	
Burundi	-3.53*	-20.89*	-3.48	
Cameroon	-1.69	-5.99	-5.20*	[1993:12]
Canada	-1.07	-2.63	-3.41	
Central African Republic	-1.93	-7.29	-5.87*	[1993:12]
Chile	-1.48	-4.32	-3.95	
Colombia	-1.60	-5.99	-3.16	
Costa Rica	-4.03*	-27.52*	-5.01*	[1998:12]
Côte d'Ivoire	-2.13	-8.56	-4.89*	[1993:12]
Dominica	-2.98	-14.64	-3.15	
Ecuador	-3.82*	-26.99*	-3.70	
Ethiopia	-1.28	-4.02	-4.65*	[1993:03]
Ghana	-2.44	-11.52	-4.87*	[1983:09]
Guatemala	-1.87	-8.26	-3.05	
Honduras	-2.12	-9.08	-3.32	
Iceland	-3.66*	-25.98*	-4.22	
India	-2.07	-8.02	-3.51	
Indonesia	-2.59	-14.13	-4.88*	[1997:10]
Kenya	-3.73*	-28.85*	-5.19*	[1995:05]
Madagascar	-2.64	-14.66	-5.29*	[1986:04]
Malawi	-3.01	-17.69	-4.66*	[1994:08]
Malaysia	-2.05	-8.16	-2.96	
Mali	-2.07	-8.47	-5.61*	[1993:12]
Mauritania	-2.47	-12.65	-5.21*	[1998:03]
Mauritius	-2.11	-8.81	-5.12*	[1986:07]
Mexico	-2.28	-11.75	-3.10	
Morocco	-2.07	-6.77	-4.63*	[1992:12]
Mozambique	-1.93	-7.23	-3.35	
Myanmar	-3.29	1.48	-3.32	
New Zealand	-2.47	-12.46	-2.70	
Nicaragua	-2.91	-16.50	-3.22	
Niger	-2.19	-9.13	-6.49*	[1993:12]
Norway	-3.03	-15.70	-4.51	
Pakistan	-2.19	-9.31	-3.88	
Papua New Guinea	-2.47	-12.38	-4.76*	[1995:03]

Cointegration Tests: Real Exchange Rate and Real Commodity Prices, Commodity-Exporting Countries, 1980-2002 (Concluded)

Country (1)	$Z(t)$ (2)	$Z(\alpha)$ (3)	$Z(t)$* (4)	Shift date (5)
Paraguay	-3.65*	-25.64*	-4.32	
Philippines	-2.94	-16.84	-3.17	
Peru	-3.17	-19.39	-6.28*	[1988:12]
Senegal	-1.59	-5.35	-5.65*	[1993:12]
South Africa	-1.68	-9.27	-2.78	
Sri Lanka	-2.51	-13.54	-4.19	
St. Vincent & Grenadines	-2.47	-11.25	-3.32	
Sudan	-2.38	-11.00	-3.51	
Suriname	-2.68	-13.91	-3.56	
Syrian Arab Republic	-1.51	-4.32	-4.80*	[1988:05]
Tanzania	-2.18	-9.75	-3.67	
Thailand	-3.25	-19.34	-3.85	
Togo	-2.50	-12.14	-5.32*	[1993:12]
Tunisia	-2.92	-16.69	-6.36*	[1986:06]
Turkey	-3.10	-17.32	-3.94	
Uganda	-3.25	-18.17	-3.92	
Uruguay	-1.77	-6.14	-3.55	
Zambia	-3.42*	-22.39*	-3.67	
Zimbabwe	-1.20	-6.24	-1.60	

Notes: The data (described in Appendices I and II) for all countries is monthly, and are expressed in logarithmic form. The estimated regression from which the residuals are derived is:
$REER = \beta_0 + \beta_1 RCOMP + \varepsilon$, where *REER* is the country's real effective exchange rate; *RCOMP* the national real commodity price; and ε is the residual. Column (2): the 5 (10) percent critical values (for T=267) for the Phillips-Ouliaris (1990) residual-based $Z(t)$ test (with a constant) are -3.36 (-3.06), based on MacKinnon (1991). Column (3): the 5 (10) percent critical value (for T=250) for the Phillips-Ouliaris (1990) residual-based $Z(\alpha)$ test (with a constant) is -20.05 (-16.65), taken from Haug (1992). Column (4): the 5 (10) percent critical value for the Gregory-Hansen (1996a) $Z(t)$* test for the presence of a level shift in the cointegrating vector is -4.61 (-4.34); the date in which the structural change is estimated to occur is given in square brackets (column (5)). For columns 2-4, an asterisk (*) denotes statistical significance at the 5 percent level, indicating that the null hypothesis of no cointegration can be rejected.

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