Journal of Development Economics 75 (2004) 239–268
JOURNAL OF Development ECONOMICS
www.elsevier.com/locate/econbase

Commodity currencies and the real exchange rate

Paul Cashin[a,*], Luis F. Céspedes[b], Ratna Sahay[a]

[a] Research Department, International Monetary Fund, 700 19th Street NW, Washington, DC 20431, USA
[b] Research Department, Central Bank of Chile, Agustinas 1180, Santiago, Chile

Received 1 January 2003; accepted 1 August 2003

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 International Monetary Fund (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 commodity-exporting countries over 1980–2002. Evidence of a long-run relationship between national real exchange rate and real commodity prices is found for about one-third of the commodity-exporting countries. The long-run real exchange rate of these 'commodity currencies' is not constant (as would be implied by purchasing power parity-based models) but is time varying, being dependent on movements in the real price of commodity exports.

JEL classification: F31; F41; O13
Keywords: Real exchange rate; Commodity prices; Developing country; Purchasing power parity

> 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).

1. Introduction

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)

* Corresponding author. Tel.: +1-202-623-6104; fax: +1-202-623-4343.
E-mail addresses: pcashin@imf.org (P. Cashin), lcespede@bcentral.cl (L.F. Céspedes), rsahay@imf.org (R. Sahay).

0304-3878/$ - see front matter
doi:10.1016/j.jdeveco.2003.08.005

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; 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.

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 Díaz-Alejandro, 1982; Edwards, 1989). There is also an extensive literature for some developed countries which links exogenous movements in the terms of trade and changes in their real exchange rates, particularly for commodity exporters Canada and Australia (see, among others, Amano and van Norden, 1995; Gruen and Wilkinson, 1994).

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). In the context of commodity-exporting countries, almost all of which are also developing countries, the real factor of primary interest in the determination of the real exchange rate is the terms of trade.

Indeed, because primary commodities dominate the exports of developing countries, fluctuations in world commodity prices have the potential to explain a large share of movements in their terms of trade. While terms of trade fluctuations have been considered a key determinant of real exchange rates (De Gregorio and Wolf, 1994; Chinn and Johnston, 1996; Montiel, 1997), it is surprising that there has been no comprehensive empirical work done to assess the mechanisms through which changes in real commodity prices affect the real exchange rate.[1] This paper takes Keynes (1930) seriously, and will examine whether price movements within the tradable sector, in particular changes in the relative price of commodity exports to imports, are a major determinant of movements in real exchange rates of commodity-dependent countries. In doing so, we are not claiming that real commodity prices have a unique role in the determination of the real exchange rate, but commodity prices are likely to be the most important source of persistent changes in the real exchange rate of commodity-dependent countries.

Importantly, Obstfeld and Rogoff (2000) point out that in the presence of sticky producer prices and perfect passthroughs, 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 passthrough 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 expressed in world prices to ameliorate 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, 2003).[2]

The key objective of this paper is to determine how many commodity-exporting countries have ‘commodity currencies’, in that movements in real commodity prices can explain fluctuations in their real exchange rates. The paper does so in several ways. First, a new monthly data set of country-specific export price indices is constructed for 58 countries over the period from 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 one-third 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 parity. Accordingly, the long-run real exchange rate of ‘commodity currencies’ is not constant (as would be implied by

[1] Two earlier country-specific analyses have been Edwards (1985), which examined the relationship between real coffee prices and Colombia’s real exchange rate; and Chen and Rogoff (2003), which found that commodity price movements influence the real exchange rates of developed country commodity-exporters Australia, New Zealand, and Canada.

[2] 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. Bidarkota and Crucini (2000) and Baxter and Kouparitsas (2000) 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.

purchasing power parity-based models) but is time varying, being dependent on movements in real commodity prices. Third, weak exogeneity tests carried out within an error-correction framework indicate that for most commodity currencies, causality runs from real commodity prices to the real exchange rate. When deviations from the long-run equilibrium relationship occur in commodity currencies, it is usually the real exchange rate that adjusts to restore long-run equilibrium. For commodity currencies, the average half-life of adjustment of the real exchange rate to its equilibrium with real commodity prices is about 10 months, which is much shorter-lived than Rogoff's (1996) consensus estimate of the half-life of real exchange rate deviations from purchasing power parity of between 3 and 5 years.

The paper is organized as follows. Section 2 briefly sets out the theoretical relationship between real commodity prices and the real exchange rate. Section 3 explains the sources and construction of the national real exchange rate and real commodity export-price data used in this study. Section 4 applies cointegration and 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 draws inferences regarding causality between the two series, and examines the speed of reversion of 'commodity currency' real exchange rates to their time-varying (commodity-price-dependent) long-run equilibrium. Section 5 concludes.

2. Theoretical framework

In describing the theoretical link between the real exchange rate and real commodity prices, we consider a small open economy that produces two different types of goods: a nontradable good and an exportable good (see Appendix A for additional details). 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).

The domestic economy is composed of two different sectors: one producing an exportable good, called "primary commodity", and the other producing a nontraded good. Firms in the export and nontraded sectors use only labor in order to produce these goods. In particular, we assume that production is carried out by competitive firms which have access to a constant returns to scale technology. Labor is free to move across sectors, thereby ensuring that wages are equated across sectors and that only supply side factors are relevant. Accordingly, we abstract from demand-side considerations and concentrate on a representation of long-run relative price determination.

Domestic consumers supply labor inelastically and consume both a nontraded and a final tradable good. This tradable good is imported from the rest of the world and is not produced domestically. Foreign firms use the primary commodity joint with an intermediate good, produced only abroad, as inputs in the production process of the final tradable good. Additionally, foreign households consume the final tradable good and a nontraded good (produced abroad).

For the purpose of the empirical analysis, we define the real exchange rate as the foreign price of the domestic basket of consumption (*EP*) relative to the foreign price of the foreign basket of consumption (P^*). After some algebra, provided in Appendix A, we can show that the determination of the real exchange rate may be summarized by the following relation:

$$\frac{EP}{P^*} = \left(\frac{a_X}{a_I^*} \frac{a_N^*}{a_N} \frac{P_X^*}{P_I^*} \right)^{\gamma} \tag{1}$$

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 source 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.

3. 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 (IMF) International Financial Statistics (IFS) and Information Notice System (INS) databases over the period from January 1980 to March 2002, which gives a total of 267 observations.

3.1. Real exchange rates and real commodity prices

The definition of the real exchange rate is the real effective exchange rate (REER) based on consumer prices. As such, we will examine the behavior of REER based on the following: (i) the nominal effective exchange rate, 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). 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: $\mathrm{REER}_i = \left[(P_iR_i)/\exp\sum_{i=1}^{n}(W_{ij}\ln(P_jR_j))\right]$ 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 1988–1990 average data on the composition of trade in manufacturing, nonoil 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 US dollars. 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). The national REER series are expressed in logarithmic form (see Appendix B for additional details).

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).[3] 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: $\mathrm{NCOMP} = \exp\{\sum_{k=1}^{K}(W_k(\ln P_k))\}$, 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).[4] Importantly, each country's NCOMP will be unique, because W_k is country specific.[5] The national RCOMP series are expressed in logarithmic form (see Appendix C 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 (Montiel, 1997) or aggregate

[3] 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 et al., 2000).

[4] In this paper, 'commodity exports' are defined as nonfuel primary product (agricultural and mineral primary products) exports—see Appendices A and B for additional details.

[5] Baxter and Kouparitsas (2000) show that, for nonfuel commodity exporters, the terms of trade are 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% of the import basket), nonfuel goods (20%) and fuels (15%). Accordingly, they find that cross-country differences in movements in the terms of trade largely emerge on the export-price side.

(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).[6] Few exporters of 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 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 et al., 2002a).[7]

3.2. 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 1988–1992, the midpoint of our sample (IMF, 1996). 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 5–8, 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 1988–1992 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, 1996).

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, eight were excluded due to the unavailability of UN COMTRADE data on their commodity exports, leaving 53 developing countries in our sample. In addition, five 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.

[6] 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.

[7] 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 a potentially exogenous variable, and so exclude volume effects of changes in commodity-export prices.

Table 1
Principal commodity exports and share of primary commodities in total exports, 1991 – 1999

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
Senegal	Phosphate rock	Groundnut oil	Fish	29	29	16	26
South Africa	Gold	Coal	Iron	46	20	5	39
Sri Lanka	Tea	Natural Rubber	Tobacco	78	9	6	20
St. Vincent and Grenadines	Bananas	Wheat	Rice	60	23	17	72
Sudan	Cotton	Gold	Sugar	45	12	12	44
Suriname	Aluminum	Rice	Nickel	80	8	5	86

Table 1 (*continued*)

Country	Principal exports			Share of exports			
	1	2	3	1	2	3	44
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.
Columns marked 1–3 denote the three largest commodity exports of each country, and their share (in %) of total commodity exports. The column marked 44 denotes the share (in %) 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. All data are averages of annual data for the period 1991–1999. See Appendix C for additional details.

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%. Among sub-Saharan African countries, commodity exports typically exceeded 50% of total exports, especially for Burundi (97%), Madagascar (90%) and Zambia (88%). Even among developed countries, the share of primary commodity exports in total exports is quite high (Australia, 54%; Iceland, 56%). In addition, many countries remain overwhelmingly dependent on export receipts from their dominant commodity exportable—cases where the dominant exportable exceeded 90% 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 four selected countries is set out in Fig. 1—Australia and Burundi (which have flexible nominal exchange rates) and Mali and Togo (which have fixed nominal exchange rates). 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 the 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 (i.e., nonstationarity in the REER). Typically, the evolution of the REER appears to be a highly persistent, slow-moving process; the REER does not appear to cycle about any particular equilibrium value. 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 (which includes Mali and Togo). Fig. 1 also displays the RCOMP indices (base 1990 = 100) for the four selected countries. Using ocular regression methods, it is readily apparent that many countries display a close relationship between

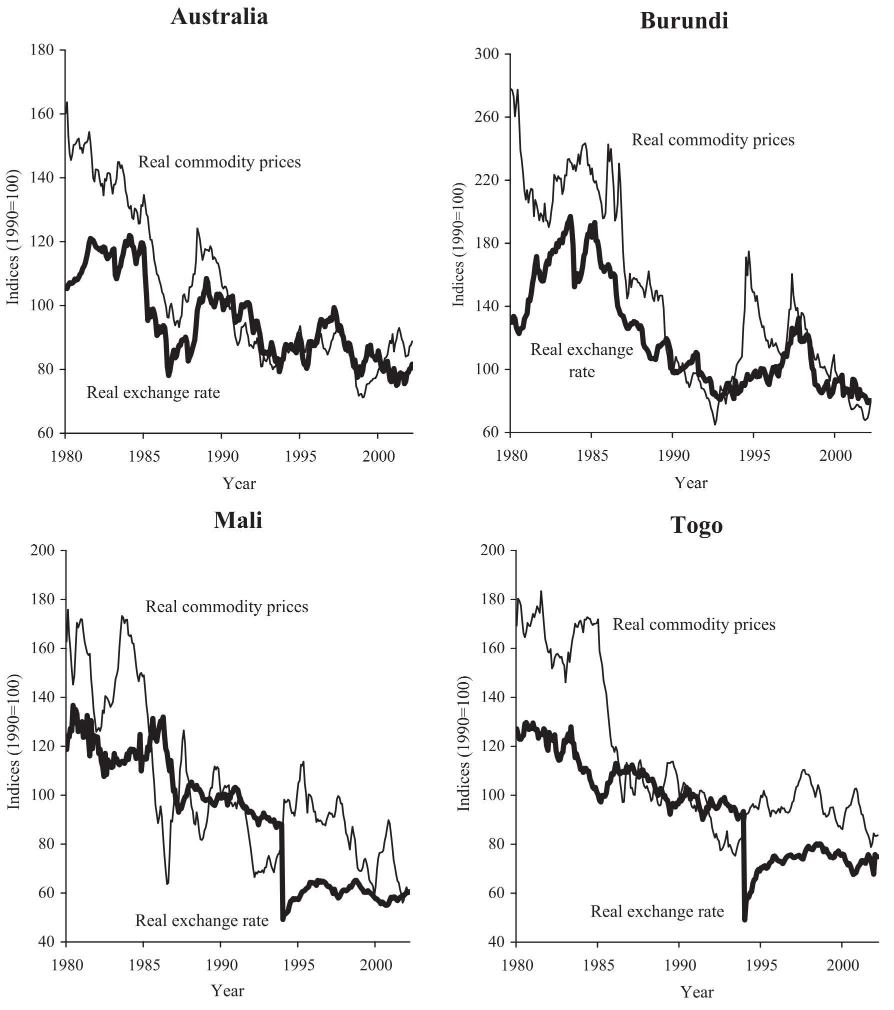

Fig. 1. Real exchange rate and real commodity prices, selected commodity-exporting countries, 1980:01 – 2002:03.

their real commodity prices and real exchange rates (such as Australia and Burundi), while others appear to display a close relationship once a one-time shift in the mean real exchange rate is accounted for (such as Mali and Togo). In the following section we will examine these relationships in some detail.

4. Empirical analysis of comovement

We use the Engle and Granger (1987) cointegration approach to assess 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. This approach allows us to examine the usefulness of specifying the real exchange rate simply as a function of real commodity prices. 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 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 with real commodity prices.

4.1. Is there a long-run relationship between 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 real exchange rate is an integrated process, then the fundamental determinants of the real exchange rate should themselves be integrated processes. In addition, nonstationarity of the real exchange rate means cointegration methods should be used to examine whether there is a long-run relationship between the fundamentals and the real exchange rate.

As set out in Section 2 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 real exchange rate and real commodity price 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.

4.1.1. Order of integration of the series

We use the Phillips and 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 prewhitened 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.[8] The results of these tests for the stationarity of the real

[8] We also applied the Zivot and 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 Mauritania and Togo, and the RCOMP series of Chile, are nonstationary in levels) these conclusions seem reasonable. The detailed results of the unit root tests are available from the authors.

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 et al., 2000).

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. An examination for the existence of cointegration is an important check on the adequacy of our model. If the long-run real exchange rate is determined by factors other than real commodity prices, then their omission should prevent us from finding evidence of cointegration. However, a finding of cointegration would imply that real commodity prices adequately capture all the permanent innovations in the real exchange rate over the sample period (Amano and van Norden, 1995).

4.1.2. Examining for cointegration: allowing for structural change

When examining data drawn from time periods characterized by changing institutional developments, the failure to find a long-run (cointegrating) relationship between a group of variables could in fact reflect a cointegrating relationship that has experienced a 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) and real commodity prices (RCOMP) 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 changes in the policy framework of 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. 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 and Hansen (1996a) test for cointegration is therefore helpful in this instance, because it allows for the timing of any regime shift to be unknown a priori.

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

$$\mathrm{REER}_i = \beta_0 + \beta_1 \mathrm{RCOMP}_t + \varepsilon_t, \qquad t = 1, \ldots, T, \tag{2}$$

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 Eq. (2), Gregory and Hansen propose a model where structural change occurs with a shift in the intercept term:

$$\mathrm{REER}_t = \beta_0 + \beta_1 \mathrm{RCOMP}_t + \beta_2 \varphi_{t\pi} + \varepsilon_t, \qquad t = 1, \ldots, T, \tag{3}$$

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:

$$\varphi_{t\pi} = \begin{cases} 0 & \text{if} \quad t \leq [T\pi] \\ 1 & \text{if} \quad t > [T\pi], \end{cases} \tag{4}$$

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 Eq. (3). In doing so, the Phillips–Perron $Z(t)$ cointegration test statistic is computed for each possible shift $\pi \in \Pi$, using the residuals from the cointegrating regression of Eq. (3). The set Π can be any compact subset of (0,1), but following Gregory and Hansen (1996a), Π is here taken to be the compact subset Π=[0.15T,0.85T]. The π is chosen so that $Z(t)$ takes the smallest value (largest negative value) across all possible break points, since the smallest $Z(t)$ gives the least favorable result for the null hypothesis (i.e., the greater chance of rejecting the null of no cointegration). We will denote the smallest of these $Z(t)$ statistics as $Z(t)$*.

While the Gregory and 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.

The Phillips and 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 and Hansen (1996a) model is also no cointegration between REER and RCOMP, 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 and Hansen (1996a) cointegration test are set out in Appendix D. 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)).[9] In addition, 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, so there is a long-run cointegrating relationship between REER and RCOMP (see Appendix D). Our finding of cointegration indicates that real commodity prices capture the permanent innovations in the real exchange rates of these 'commodity currencies'.

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 Fully Modified (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.[10] Using the Hansen (1992) tests we find no evidence of an unstable relationship between REER and RCOMP for any of the above three countries, and so conclude that there is cointegration with stable parameters.

4.1.3. 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

[9] This concept of a long-run relationship which is subject to structural change formalizes the idea of Dornbusch and Vogelsang (1991) of PPP holding once allowance is made for a shift in the mean level of the real exchange rate. Work by Flynn and Boucher (1993) and Hegwood and Papell (1998) also allows for a structural break in cointegration analyses of the determinants of the real exchange rate. Flynn and Boucher (1993) find that cointegration analyses are biased against finding stationarity of the residuals from the long-run regression if allowance is not made for structural breaks [typically caused by government interventions affecting the level of the nominal (and real) exchange rate].

[10] 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 $[0.15T, 0.85T]$. Hansen's (1992) parameter stability tests are based on the residuals of a Fully Modified least-squares regression (Phillips and Hansen, 1990).

RCOMP [as set out in Eq. (3)] 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.[11] 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.[12,13] Estimates of the commodity price elasticity of the real exchange rate are statistically significant, 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 10 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 Eq. (2)] was again estimated using Phillips and Hansen's (1990) FM method—the results are set out in the upper panel of Table 2.[14] 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

[11] For the Phillips and Hansen (1990) FM estimation, we employ the Bartlett kernel, Andrews' (1991) automatic bandwidth selector and the prewhitened 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.

[12] 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.

[13] While the null hypothesis of no cointegration could be rejected in favor of the alternative hypothesis of cointegration (allowing for a structural shift) for Cameroon, Côte d'Ivoire, Ethiopia, Madagascar, Mauritius, Peru, Syrian Arab Republic and Senegal, for these countries the coefficient on RCOMP in the cointegrating regression was found to be either negative or positive (yet not 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 Table 3.

[14] 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 Table 3.

Table 2
Cointegration and stability tests, real exchange rate and real commodity prices, 1980–2002

Country (1)	Cointegrating parameters		R^2 (3)	Hansen tests		
	RCOMP (2a)	DUM (2b)		Lc (4)	MeanF (5)	SupF (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						
Central African Republic	0.230 (0.058)	– 0.506 (0.034)	0.909			
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			
Togo	0.297 (0.059)	– 0.308 (0.030)	0.868			
Tunisia	0.164 (0.061)	– 0.291 (0.024)	0.964			

The data (described in Appendices A and B) 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): $\mathrm{REER}=\beta_0+\beta_1\mathrm{RCOMP}+\beta_2\mathrm{DUM}+\varepsilon$, 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 prewhitened 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) 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. Gregory and Hansen (1996a) tests for the presence of a regime shift in the cointegrating vector (reported in Appendix D) reveal that the null hypothesis of no cointegration can be rejected, indicating a significant level shift in the cointegrating relation for the following: 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).

indicate that there is no evidence of instability in the relationship between each country's REER and the RCOMP (at the 5% 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.

For those ‘commodity currency’ countries that indicate that there is a long-run relationship between each country’s REER and RCOMP (the eight countries exhibiting cointegration with stable parameters and the 11 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 0.162 (for Iceland) to 2.03 (for Ecuador). Across all commodity currencies, the median value of the elasticity is 0.42, indicating that a 10% rise in real commodity prices is typically associated with a 4.2% appreciation of the real exchange rate.

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 19 countries, over 85% 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 remarkably strong result. Clearly, movements in real commodity prices are an important driver of real exchange rates in such commodity-dependent countries.[15]

In summary, standard cointegration tests provide evidence of long-run relationships between the real exchange rate and real commodity prices. The evidence for a cointegrating relationship between these variables, allowing for a structural shift (of unknown timing), is also 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 typically from nominal exchange rate devaluation.[16,17] For one-third (19 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 dependent on the evolution of world real commodity prices.

[15] Cointegration of two or more variables implies that these variables move together over time such that they revert to a long-run equilibrium relationship. Given that real commodity prices and the real exchange rate are cointegrated (as was found for 19 ‘commodity currencies’), then the econometric problems typically associated with exogeneity issues (such as simultaneity bias, consistency and identification) are asymptotically negligible in such static cointegrating regressions (see McDermott and Wong, 1990). Indeed, if a set of nonstationary variables are cointegrated (as defined by Engle and Granger, 1987), then the concept of exogeneity is not useful, as a particular attraction of cointegrating regressions is that all of the variables may be treated as jointly endogenous.

[16] In many countries (and especially those with pegged nominal exchange rates), real exchange rate movements occur chiefly through large and rapid nominal devaluation, rather than through cumulative inflation differentials (Goldfajn and Valdés, 1999). In addition, and as pointed out by a referee, if the speed of adjustment of the real exchange rate to its equilibrium is very slow, a long period of exchange rate misalignment that is abruptly corrected by a devaluation may be being picked up in the estimation as a change in the constant term of the cointegrating regression.

[17] In the case of commodity currencies, the real exchange rate has a time-varying long-run relationship with real commodity prices, after taking into account the structural shift in the cointegrating relationship induced by a level shift in the real exchange rate. Moreover, just as univariate unit root tests for PPP-based equilibrium real exchange rates will be biased toward nonrejection of the unit root null if potential structural breaks (intercept changes) are not accounted for (Perron and Vogelsang, 1992; Papell, 2002), cointegration tests of whether the real exchange rate has a time-varying equilibrium relationship with its real fundamentals will also be biased toward nonrejection of the no cointegration null if potential structural breaks are not accounted for (Gregory and Hansen, 1996a). Indeed, the failure of empirical models of exchange rate determination to account for structural breaks in the long-run relationship between the real exchange rate and its fundamentals could account for a sizeable share of the poor empirical performance of such models.

4.2. Causality and exogeneity tests

Evidence of cointegration rules out the possibility of the estimated relationship being a "spurious regression". As noted in Section 4.1, for about one-third 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 the Granger sense) in *at least* one direction, and the direction of causality can be ascertained using the error correction methodology suggested by Engle 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. Prior to estimating the error-correction model, we follow Engle (1984) and Engle et al. (1983) and apply a Lagrange Multiplier (LM) test statistic, to test for weak exogeneity of real commodity prices in the real exchange rate error-correction equation. In addition, a Likelihood Ratio (LR) test statistic is applied to the joint significance of the sum of the lags of each explanatory variable, to test for strict or "short-term" Granger noncausality.[18]

The weak exogeneity and Granger noncausality tests are conducted using the error-correction procedure, only for those countries where a cointegrating relationship between the real commodity price (RCOMP) and real exchange rate (REER) has been established. In error correction form the model becomes:

$$\Delta \mathrm{REER}_t = \eta + \sum_{i=1}^{p-1} \alpha_i \Delta \mathrm{REER}_{t-i} + \sum_{j=1}^{p-1} \beta_j \Delta \mathrm{RCOMP}_{t-j} + \Theta(\mathrm{REER} - \kappa \mathrm{RCOMP})_{t-1} + e_t \tag{5}$$

$$\Delta \mathrm{RCOMP}_t = \eta' + \sum_{i=1}^{p-1} \gamma_i \Delta \mathrm{REER}_{t-i} + \sum_{j=1}^{p-1} \delta_j \Delta \mathrm{RCOMP}_{t-j} + \Omega(\mathrm{REER} - \kappa \mathrm{RCOMP})_{t-1} + e_t' \tag{6}$$

where η and η' are constant terms; e and e' are disturbance terms; and the lagged error-correction term $(\mathrm{REER} - \kappa\mathrm{RCOMP})_{t-1}$ is the lagged residual from the cointegrat-

[18] Real commodity prices are weakly exogenous with respect to the real exchange rate if inference can be conducted conditional on the sample values of real commodity prices with no loss of relevant sample information (see Engle et al., 1983). Rejection of the null hypothesis of Granger noncausality implies that one variable can be predicted using past values of another variable; that is, real commodity prices Granger cause the real exchange rate if the real exchange rate can be predicted from past values of real commodity prices. In an error-correction model, one or more of the differenced variables must be Granger-caused by the lagged error-correction term.

ing regression [of Eqs. (2) and (3)] between REER_t and RCOMP_t, and measures the deviation from purchasing power parity in the previous period.[19] In Eq. (5), REER_t is influenced by RCOMP_t either through the lagged dynamic terms of RCOMP_t if all the β_j are not equal to zero ('short-run' Granger causality), or by the lagged error-correction term if Θ is nonzero ('long-run' Granger causality). The speed of adjustment parameters [Θ and Ω in Eqs. (5) and (6)] indicates how quickly the system returns to its long-run equilibrium after a temporary departure from it.

The null hypothesis of Engle's (1984) weak exogeneity test is $\Omega = \sigma_{12} = 0$, where $\sigma_{12} = \mathrm{corr}(e_t, e'_t)$.[20] Nonrejection of the null of weak exogeneity of real commodity prices implies that real commodity prices are exogenous to the system and do not respond to any deviation from long-run equilibrium, and accordingly that all of the adjustment to deviations from the long-run equilibrium (through the error-correction component) correspond to adjustments in the real exchange rate. That is, nonrejection of the weak exogeneity null implies that Eq. (6) is redundant.

For those countries with commodity currencies, the results of the causality and exogeneity analysis using the error-correction procedure are set out in Table 3. The weak exogeneity test results support the hypothesis that real commodity prices are statistically exogenous for 10 of the 19 commodity-currency countries (at the 1% level of significance). The resulting TR^2 is small for these 10 commodity currencies, indicating that the data does not contain any evidence against the hypothesis of weak exogeneity of real commodity prices, and that the disequilibrium error from the cointegrating relationship significantly influences changes in the real exchange rate. This result implies that for these countries, the real exchange rate–real commodity price relationship can be modeled in a single-equation error-correction framework.[21] For the 10 countries satisfying Engle's (1984) test of weak exogeneity of real commodity prices, with Θ less than zero (which ensures error correction) and statistically significant, a positive (negative) disequilibrium term $(\mathrm{REER}-\kappa\mathrm{RCOMP})_{t-1}$ will ensure that REER declines (rises) toward its long-run equilibrium path.[22] These

[19] 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 4.1.2 indicated Eq. (3) was the appropriate cointegrating regression.

[20] We are grateful to an anonymous referee for suggesting further analysis of weak exogeneity issues. The appropriate test of the weak exogeneity of RCOMP is done by testing $\tilde{u}_1$ and $[(\mathrm{REER}-\kappa\mathrm{COMP})_{t-1}-\hat{\Theta}\tilde{u}_2]$ as omitted from Eq. (6), where $\hat{\Theta}$ is the estimated parameter on the error-correction term in Eq. (5), $\tilde{u}_1$ are the residuals from Eq. (5) and $\tilde{u}_2$ are the residuals from the regression of ΔRCOMP on η', $\Sigma\Delta$REER and $\Sigma\Delta$RCOMP. The test of weak exogeneity of real commodity prices is computed as $TR^2(T=265)$ of the regression of $\tilde{u}_2$ on $[(\mathrm{REER}-\kappa\mathrm{COMP})_{t-1}-\hat{\Theta}\,\tilde{u}_2]$, $\tilde{u}_1$, $\Sigma\Delta$REER and $\Sigma\Delta$RCOMP. This statistic is asymptotically distributed as $\chi^2(2)$ under the null; the 5% (1%) critical value is 5.99 (9.21). The weak exogeneity test is biased toward rejecting the null hypothesis of weak exogeneity if there exists any form of misspecification in the model. For additional details on the LM test of weak exogeneity, see Engle (1984) and McDermott and Wong (1990).

[21] In addition, for five countries (such as Bangladesh and Togo), the LR test indicates there is evidence that short-run movements in RCOMP help predict (Granger cause) part of the short-run movement in REER.

[22] 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.

Table 3
Real exchange rate (REER) and real commodity prices (RCOMP): exogeneity and causality, 1980–2002

Country (1)	Half-life of real exchange rate deviations from PPP (months) (2)	Half-life of real exchange rate deviations from commodity-price equilibrium (months) (3)	Lag (4)	Engel test of weak exogeneity: χ^2-statistic (5)	$\beta_j = 0$: χ^2-statistic (p-value) (6)	Θ (t-statistic) (7)
Bangladesh	20.66	12.25	1	1.97	15.49 (0.00)	− 0.055 (− 1.78)
Bolivia	5.73	2.38	1	1.03	0.71 (0.70)	− 0.253 (− 5.96)
Burundi	76.67	11.60	1	4.97	3.12 (0.21)	− 0.058 (− 3.31)
Central African Republic	86.30	11.81	1	4.83	15.05 (0.00)	− 0.057 (− 3.91)
Ghana	66.95	9.99	1	6.82	3.46 (0.17)	− 0.067 (− 3.33)
Iceland	97.28	8.10	1	4.15	4.37 (0.11)	− 0.082 (− 3.55)
Kenya	20.66	7.44	1	5.54	28.06 (0.00)	− 0.089 (− 3.86)
Mali	76.67	9.41	1	1.04	24.24 (0.00)	− 0.071 (− 4.13)
Papua New Guinea	20.53	13.24	2	2.31	2.14 (0.34)	− 0.051 (− 1.91)
Togo	27.27	10.15	1	6.09	28.11 (0.00)	− 0.066 (− 4.04)
Median	39.72	10.07				− 0.067
Standard deviation						0.060

See the error-correction model of Eq. (5). 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 4.1 indicated was appropriate. Column 2: 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 (HL) is calculated using the formula: HL = ABS(log(1/2)/log(α)), where α is the autoregressive parameter derived from the Dickey–Fuller [or AR(1)] least-squares regression of the real exchange rate. Column 3: The implied half-life of real exchange rate deviations from commodity-price equilibrium is calculated as follows. The time (T) required to dissipate x percent (in this case, 50%) of the deviation is determined according to: $(1-\Theta)^T=(1-x)$, where Θ is the coefficient of the error-correction term (given in column 7) and T is the required number of periods (months). Lagrange Multiplier tests for serial correlation (with the order of serial correlation tested being one more than the optimal lag length of the error-correction model) indicate that there is little evidence of residual serial correlation. Column 4: Lag is the number of lagged first-difference terms in the error-correction model of Eq. (5), and is determined by minimizing the Akaike Information criterion. Column 5: The critical value of Engel's (1984) Lagrange Multiplier test of weak exogeneity is distributed as a χ^2 (2); the value is 5.99 (9.21) at the 5% (1%) level of significance. Test statistics less than the critical value indicate that the null hypothesis of weak exogeneity of RCOMP in the REER error-correction model of Eq. (5) cannot be rejected. Column 6: The value of the Likelihood Ratio test statistic of the null hypothesis that all the $\beta_j = 0$ in the error-correction model of Eq. (5), with associated p-value in parentheses. Column 7: The value of the parameter on the error-correction term (Θ) in Eq. (5), with associated t-statistic in parentheses.

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 for the majority of commodity currencies it is solely the real exchange rate which adjusts to preserve the long-run equilibrium with commodity prices, and there is evidence in support of the notion of rising real commodity prices leading to increasing (appreciating) real exchange rates.[23]

4.3. Commodity currencies and the PPP puzzle

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 proposition that exchange rates are determined by movements in relative prices. 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%) to the real exchange rate is about 3–5 years, implying a slow speed of reversion to (constant) parity of between 13% and 20% 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 1–2 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, 2003). 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 et al., 2000, 2002a; 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-dependent 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

[23] Nine countries rejected the null hypothesis of weak exogeneity of real commodity prices—in all cases (except Australia) the null was rejected not because the error terms in Eqs. (5) and (6) were correlated (i.e., not because corr(e_t, e_t') $\neq$ 0 in Engle's (1984) exogeneity test), but because of the endogeneity of RCOMP {that is, RCOMP is adjusting to restore the long-run equilibrium relationship with REER [$\Omega \neq 0$ in Eq. (6)]}. This finding is of some comfort, as correlation of the error terms in Eqs. (5) and (6) would imply that the error-correction models were misspecified. Importantly, in all nine countries rejecting weak exogeneity of real commodity prices the estimated Θ in Eq. (5) was negative (which ensures error correction) and statistically significant. Accordingly, for these nine countries both the real exchange rate and real commodity prices adjust to close any given deviation from long-run equilibrium, so that the speed of reversion of REER to deviations from long-run equilibrium cannot be calculated for these commodity currencies.

regression), without controlling for commodity prices, and focus on the magnitude of the least squares estimates of the autoregressive parameter. Across all countries, the median half-life of parity reversion is 36 months for our sample of 58 commodity-dependent countries, while for the 19 'commodity currencies' the median half-life of parity reversion is somewhat longer at 49 months.[24] These results are consistent with Rogoff's (1996) consensus of half-lives of parity reversion of between 36 and 60 months (3–5 years).

Next we turn to the results from our error-correction model, which provides information on the speed with which real exchange rates adjust to reestablish their long-run equilibrium relationship with real commodity prices (see column 3 of Table 3). The magnitude of Θ [the coefficient on the error-correction term in Eq. (5)] indicates that for some countries (such as Bangladesh and Papua New Guinea) only about 5% of the deviation of the REER from long-run equilibrium is eliminated in 1 month (implying a half-life of parity deviation of about 13 months), while for other countries (such as Kenya and Iceland) about 8% of the deviation is eliminated in 1 month (implying a half-life of parity deviation of about 8 months), a very rapid speed of adjustment. For each of the 10 'commodity currency' countries with weakly exogenous commodity prices, the half-life of the reversion of the real exchange rate to its (constant) long-run *average* level (reported in column 2 of Table 3) is much longer than the half-life of the reversion of the real exchange rate to its (time varying) long-run *equilibrium* with real commodity prices (reported in column 3 of Table 3).[25]

Averaging across these 10 'commodity currency' countries, the median speed of error correction on real exchange rates is about 6.5% per month (Table 3). The elimination of 6.5% 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 10 months, which is much smaller than the typical half-life (of about 3–5 years) reported in the simple PPP-based regressions analyzed above (Rogoff, 1996). That is, while the real exchange rate of these 10 countries has a slow reversion to its *average* level (the median half-life of parity deviations is 39 months), it has a much faster speed of adjustment towards its long-run *equilibrium* (the median half-life of deviations is about 10 months), where that equilibrium depends on the evolution of real commodity prices as a fundamental determinant of the real exchange rate (see columns 2 and 3 of Table 3). 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.

---

[24] 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 (HL) is calculated using the formula: HL = ABS(log(1/2)/log(α)), where α is the autoregressive parameter derived from the Dickey–Fuller [or AR(1)] least-squares regression. These half-life results are comparable to those obtained by Cheung and Lai (2000) using least-squares estimation on monthly bilateral (post-Bretton Woods) dollar real exchange rates for developed countries, which yielded an average half-life of 3.3 years.

[25] 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 et al., 1999).

5. 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 one-third of the commodity-exporting countries we find robust evidence in support of the long-run comovement of national real exchange rate and real commodity-export price series. While the real exchange rate and real commodity prices of these 'commodity currencies' will be subject to transitory deviations from their long-run equilibrium, these two series move together over time such that they revert to an equilibrium relationship. In addition, weak exogeneity tests indicate that, for the majority of commodity currencies, it is the real exchange rate which adjusts to restore the long-run equilibrium with real commodity prices. This group of commodity currencies is found to exhibit extremely rapid half-lives of adjustment of real exchange rates to equilibrium of about 10 months. These estimates 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 3–5 years. As presciently conjectured by Keynes (1930), purchasing power parity is a weak model of the long-run real exchange rate in countries with commodity currencies, as these countries typically experience large and long-lived real shocks. The long-run real exchange rate of commodity currencies is not constant (as would be implied by parity-based models) but is time varying, being dependent on movements in real commodity prices.

Acknowledgements

The authors are grateful to John McDermott, Paul Nicholson, Sam Ouliaris, Paolo Pesenti, Carmen Reinhart, Ken Rogoff, Miguel Savastano, seminar participants at the Third Annual IMF Research Conference, and especially a coeditor and an anonymous referee for comments and suggestions on earlier drafts. The views expressed are those of the authors, and do not necessarily represent those of the International Monetary Fund or the Central Bank of Chile.

Appendix A. Details of the theoretical framework

In this appendix we present in detail the theoretical framework contained in Section 2 of the paper. We study a small open economy that produces two types of goods, a nontraded good and an exportable good which is associated with the production of a primary commodity. The details of the model are as follows.

A.1. 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 given by $y_X = a_X L_X$, 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: $y_N = a_N L_N$, 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 (w) must be the same across sectors. Profit maximization in both sectors yields the familiar conditions: $P_X = w/a_X$ and $P_N = w/a_N$, where P_N is the price of the nontraded good and P_X the price of the primary commodity.

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:

$$P_N = \frac{a_X}{a_N} P_X \tag{A1}$$

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. Given our freely mobile labor assumption, wages and prices will also rise in the nontraded sector.

A.2. 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: $C = \kappa C_N^{\gamma} C_T^{1-\gamma}$, 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:

$$P = P_N^{\gamma} P_T^{1-\gamma} \tag{A2}$$

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. The law of one price is assumed to hold for the imported good, so that: $P_T = P_T^{*}/E$, 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.

A.3. 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.[26] 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: $Y_{\mathrm{N}}^{*}=a_{\mathrm{N}}^{*}L_{\mathrm{N}}^{*}$. 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: $Y_{\mathrm{I}}^{*}=a_{\mathrm{I}}^{*}L_{\mathrm{I}}^{*}$. Labor mobility across (foreign) sectors ensures that the (foreign) wage is equated across sectors.[27] 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:

$$P_{\mathrm{N}}^{*}=\frac{a_{\mathrm{I}}^{*}}{a_{\mathrm{N}}^{*}}P_{\mathrm{I}}^{*} \tag{A3}$$

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: $Y_{\mathrm{T}}^{*}=v(Y_{\mathrm{I}}^{*})^{\beta}(Y_{\mathrm{X}}^{*})^{1-\beta}$. It is straightforward to show that the cost of one unit of the tradable good in terms of the foreign currency is given by: $P_{\mathrm{T}}^{*}=(P_{\mathrm{I}}^{*})^{\beta}(P_{\mathrm{X}}^{*})^{1-\beta}$. Foreign consumers are assumed to consume the foreign nontraded good and this final good in the same fashion as domestic consumers. Foreigners also supply labor inelastically to the different sectors. Therefore, the consumer price index for the foreign economy can be represented by:

$$P^{*}=(P_{\mathrm{N}}^{*})^{\gamma}(P_{\mathrm{T}}^{*})^{1-\gamma}. \tag{A4}$$

A.4. 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 foreign price of the

[26] 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.

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

domestic basket of consumption relative to the foreign price of a foreign basket of consumption (EP/P^*). Using Eqs. (A1–4) we can show that:

$$\frac{EP}{P^*} = \left(\frac{a_X}{a_I^*} \frac{a_N^*}{a_N} \frac{P_X^*}{P_I^*} \right)^{\gamma} \tag{A5}$$

which is Eq. (1) as set out in Section 2 of the paper.

Appendix B. 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 D. The primary data sources are the IMF's International Financial Statistics (IFS) and Information Notice System (INS). 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 (1991–1999 average) share of each commodity in total primary commodity exports, and the IMF's (US dollar-based) data on world commodity prices (taken from the IMF's IFS). The derivation of this index is described in detail in Appendix C.

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 US 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.

Appendix C. 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 1991–1999 average total value of primary commodity exports; the 44 individual nonfuel commodity weights are calculated by dividing the 1991–1999 average value of each individual commodity export by the 1991–1999 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 (US 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 (US 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.

C.1. Nominal commodity prices

The prices (taken from the IMF's IFS) of the 44 nonfuel commodities used in the calculation of the national commodity-price indices are the following: 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. The source, and a brief description of the individual commodity prices, is available in a longer working paper version of this paper (see Cashin et al., 2002b).

Appendix D. 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	

(continued on next page)

Appendix D (*continued*)

Country (1)	$Z(t)$ (2)	$Z(\alpha)$ (3)	$Z(t)$* (4)	Shift date (5)
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]
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 and 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	

The data (described in Appendices A and B) for all countries are 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%) critical values (for T=267) for the Phillips and Ouliaris (1990) residual-based $Z(t)$ test (with a constant) are −3.36 (−3.06), based on MacKinnon (1991). Column 3: The 5% (10%) critical value (for T=250) for the Phillips and Ouliaris (1990) residual-based $Z(\alpha)$ test (with a constant) is −20.05 (−16.65), taken from Haug (1992). Column 4: The 5% (10%) critical value for the Gregory and 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% level, indicating that the null hypothesis of no cointegration can be rejected.

References

Amano, R., van Norden, S., 1995. Terms of trade and real exchange rates: the Canadian evidence. Journal of International Money and Finance 14, 83–104.

Andrews, D.W.K., 1991. Heteroskedasticity and autocorrelation consistent covariance matrix estimation. Econometrica 59, 817–858.

Andrews, D.W.K., Monahan, J.C., 1992. An improved heteroskedasticity and autocorrelation consistent covariance matrix estimation. Econometrica 60, 953–966.

Baffes, J., Elbadawi, I., O'Connell, S., 1999. Single equation estimation of the real exchange rate. In: Hinkle, L., Montiel, P. (Eds.), Exchange Rate Misalignment: Concepts and Measurement for Developing Countries. Oxford Univ. Press, Oxford, pp. 405–464.

Baxter, M., Kouparitsas, M., 2000. What causes fluctuations in the terms of trade? Working Paper 7462. National Bureau of Economic Research, Cambridge, MA.

Bidarkota, P., Crucini, M., 2000. Commodity prices and the terms of trade. Review of International Economics 8, 647–666.

Boyd, D., Smith, R., 1999. Testing for purchasing power parity: econometric issues and an application to developing countries. Manchester School 67, 287–303.

Cashin, P., McDermott, C.J., 2002. The long-run behavior of commodity prices: small trends and big variability. IMF Staff Papers 49, 175–199.

Cashin, P., Liang, H., McDermott, C.J., 2000. How persistent are shocks to world commodity prices? IMF Staff Papers 47, 177–217.

Cashin, P., McDermott, C.J., Scott, A., 2002a. Booms and slumps in world commodity prices. Journal of Development Economics 69, 277–296.

Cashin, P., Céspedes, L., Sahay, R., 2002b. Keynes, cocoa and copper: in search of commodity currencies, IMF Working Paper 02/223. International Monetary Fund, Washington, DC.

Chen, Y., Rogoff, K., 2003. Commodity currencies and empirical exchange rate puzzles. Journal of International Economics 60, 133–160.

Cheung, Y.-W., Lai, K., 2000. On the purchasing power parity puzzle. Journal of International Economics 52, 321–330.

Chinn, M., Johnston, L., 1996. Real exchange rate levels, productivity and demand shocks: evidence from a panel of 14 countries, Working Paper 5907. National Bureau of Economic Research, Cambridge, MA.

Cuddington, J., Urzua, C.M., 1989. Trends and cycles in the net barter terms of trade: a new approach. Economic Journal 99, 426–442.

De Gregorio, J., Wolf, H., 1994. Terms of trade, productivity and the real exchange rate, Working Paper 4807. National Bureau of Economic Research, Cambridge, MA.

Deaton, A., Laroque, G., 1992. On the behavior of commodity prices. Review of Economic Studies 59, 1–23.

Deaton, A., Miller, R., 1996. International commodity prices, macroeconomic performance and politics in sub-Saharan Africa. Journal of African Economies 5, 99–191 (AERC Supplement).

Dehn, J., 2000. The effects on growth of commodity price uncertainty and shocks. mimeo. Centre for the Studies of African Economies. University of Oxford, Oxford.

Díaz-Alejandro, C., 1982. Exchange rates and the terms of trade in the Argentine Republic, 1913–1976. In: Syrquin, M., Teitel, S. (Eds.), Trade, Stability, Technology, and Equity in Latin America. Academic Press, New York, pp. 27–41.

Dornbusch, R., Vogelsang, T., 1991. Real exchange rates and purchasing power parity. In: de Melo, J., Sapir, A. (Eds.), Trade Theory and Economic Reform: North, South and East, Essays in Honor of Bela Balassa. Basil Blackwell, Cambridge, MA, pp. 3–24.

Edwards, S., 1985. Commodity export prices and the real exchange rate in developing countries: coffee in Colombia, Working Paper 1570. National Bureau of Economic Research, Cambridge, MA.

Edwards, S., 1989. Real Exchange Rates, Devaluation and Adjustment: Exchange Rate Policy in Developing Countries. MIT Press, Cambridge, MA.

Engle, R.F., 1984. Wald, likelihood ratio, and Lagrange multiplier tests in econometrics. In: Griliches, Z., Intriligator, M.D. (Eds.), Handbook of Econometrics, vol. II. Elsevier, Amsterdam, pp. 776–826.

Engle, R.F., Granger, C.W.J., 1987. Cointegration and error correction: representation, estimation and testing. Econometrica 55, 251–276.

Engle, R.F., Hendry, D.F., Richard, A., 1983. Exogeneity. Econometrica 51, 227–304.

Flynn, N., Boucher, J., 1993. Tests of long-run purchasing power parity using alternative methodologies. Journal of Macroeconomics 15, 109–122.

Froot, K., Rogoff, K., 1995. Perspectives on PPP and long-run real exchange rates. In: Grossman, G., Rogoff, K. (Eds.), Handbook of International Economics, vol. 3. Elsevier, Amsterdam, pp. 1647–1688.

Goldfajn, I., Valdés, R., 1999. The aftermath of appreciations. Quarterly Journal of Economics 114, 229–262.

Gregory, A., Hansen, B., 1996a. Residual-based tests for cointegration in models with regime shifts. Journal of Econometrics 70, 99–126.

Gregory, A., Hansen, B., 1996b. Tests for cointegration in models with regime and trend shifts. Oxford Bulletin of Economics and Statistics 58, 555–560.

Grilli, E., Yang, M.-C., 1988. Primary commodity prices, manufactured goods prices, and the terms of trade in developing countries: what the long run shows. World Bank Economic Review 2, 1–47.

Gruen, D., Wilkinson, J., 1994. Australia's real exchange rate—is it explained by the terms of trade or by real interest differentials? Economic Record 70, 204–219.

Hansen, B., 1992. Tests for parameter instability in regressions with $I(1)$ processes. Journal of Business and Economic Statistics 10, 321–335.

Haug, A., 1992. Critical values for the Z_α Phillips–Ouliaris test for cointegration. Oxford Bulletin of Economics and Statistics 54, 345–351.

Hegwood, N., Papell, D., 1998. Quasi purchasing power parity. International Journal of Finance and Economics 3, 279–289.

International Monetary Fund, 1996. World Economic Outlook, International Monetary Fund, Washington DC.

Keynes, J.M., 1930. A Treatise on Money: Volume I. The Pure Theory of Money. MacMillan, London.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P., Shin, Y., 1992. Testing the null hypothesis of stationarity against the alternative of a unit root: how sure are we that economic time series have a unit root? Journal of Econometrics 54, 159–178.

MacKinnon, J., 1991. Critical values for cointegration tests. In: Engle, R., Granger, C. (Eds.), Long Run Economic Relationships: Readings in Cointegration. Oxford Univ. Press, Oxford, pp. 267–276.

McDermott, C.J., Wong, A.Y.-T., 1990. Exogeneity in the New Zealand money–income relationship. New Zealand Economic Papers 24, 24–41.

Meese, R.A., Rogoff, K., 1983. Empirical exchange rate models of the seventies: do they fit out of sample. International Economic Review 14, 3–24.

Montiel, P., 1997. Exchange rate policy and macroeconomic management in ASEAN countries. In: Hinklin, J., Robinson, D., Singh, A. (Eds.), Macroeconomic Issues Facing ASEAN Countries. International Monetary Fund, Washington, DC, pp. 253–298.

Obstfeld, M., Rogoff, K., 1996. Foundations of International Macroeconomics. MIT Press, Cambridge, MA.

Obstfeld, M., Rogoff, K., 2000. New directions for stochastic open economy models. Journal of International Economics 50, 117–153.

Papell, D., 2002. The great appreciation, the great depreciation, and the purchasing power parity hypothesis. Journal of International Economics 57, 51–82.

Perron, P., Vogelsang, T., 1992. Nonstationarity and level shifts with an application to purchasing power parity. Journal of Business and Economic Statistics 10, 301–320.

Phillips, P.C.B., Hansen, B., 1990. Statistical inference in instrumental variable regression with $I(1)$ processes. Review of Economic Studies 57, 99–125.

Phillips, P.C.B., Ouliaris, S., 1990. Asymptotic properties of residual-based tests for cointegration. Econometrica 58, 165–193.

Phillips, P.C.B., Perron, P., 1988. Testing for a unit root in time series regression. Biometrika 75, 335–346.

Reinhart, C., Rogoff, K., 2002. The modern history of exchange rate arrangements: a reinterpretation, Working Paper 8963. National Bureau of Economic Research, Cambridge, MA.

Rogoff, K., 1996. The purchasing power parity puzzle. Journal of Economic Literature 34, 647–668.

Zivot, E., Andrews, D.W.K., 1992. Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis. Journal of Business and Economic Statistics 10, 251–270.