What Makes a Commodity Currency?*

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This draft: October 10, 2013

ABSTRACT

The "*commodity currency*" literature highlights the robust exchange rate response to fluctuations in world commodity prices that occurs for major commodity exporters. The magnitude of this response, however, varies widely among countries. Our empirical analysis finds that, in accordance with theory, the long-run relationship between the real exchange rate and commodity prices depends on the nation's export market structure, its monetary policy choices and its degree of trade and financial openness. We also show that, in the short-run, the commodity price-exchange rate connection is much weaker for our large set of economies than has been observed in prior literature based on a small set of advanced economies. Given concerns for the Dutch disease or resource curse that operate through the real exchange rate, our findings are of particular relevance for monetary policy-making and for globalization strategy in commodity-exporting developing economies.

Keywords: Real exchange rate; Commodity prices; Panel cointegration; Commodity exports; Monetary policy; Globalization

* We are grateful to Stephen Turnovsky, Seik Kim, Sebastian Fossati, Kar-yiu Wong, Chang-Jin Kim, Eric Zivot and seminar participants at the University of California, Riverside, for their helpful comments and suggestions. We also thank Paul Cashin and Susan Yang at the IMF for providing some of the data used in this paper.
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1 Introduction

About a third of the countries in the world rely on primary commodities such as mineral, agricultural, and energy products as a significant source of export earnings. The wild fluctuations of global commodity prices thus account for a large share of these countries' terms-of-trade shocks, which can have a major influence on the value of their currencies. The "commodity currency" literature demonstrates the strong and robust real exchange rate response to global commodity price fluctuations and emphasizes transmission mechanisms such as terms-of-trade adjustment, the income effect, and the portfolio balance channel.[1] While an increase in the world prices of primary commodities brings about higher export revenue for their exporters, an induced corresponding real currency appreciation can crowd out the exports of non-commodity industries by undermining their price competitiveness in the world market. This so-called "Dutch Disease" consideration underscores the importance of understanding the exchange rate response to world commodity price movements as it may inform strategies for growth and policy decisions.[2]

While the literature emphasizes a generally robust exchange rate response to commodity price movements, especially for commodity exporters with a floating nominal exchange rate, little attention is paid to the wide range of response magnitudes and the reasons behind it.[3] This paper seeks to understand this variation by examining the role of potential economic determinants. Specifically, we attempt to answer the following questions: are the observed diverse commodity price elasticities across countries attributable to differences in: i) export market structure, ii) policy choices, specifically monetary, or iii) the extent of trade and financial

[1] Currencies that respond significantly to the world prices of their corresponding country's commodity exports are called "*commodity currencies*". See Amano and van Norden (1995), Chen and Rogoff (2003, 2012), MacDonald and Ricci (2004) and Cashin et al. (2004) for empirical exploration covering a range of developed and developing countries.

[2] See Corden and Neary (1982) for the core model of the Dutch disease. Using the model characterized by a non-traded good (services) and two traded goods (energy and manufactures), they address the effects of a boom in the energy sector on the distribution of income, and on the size and profitability of the manufacturing sector. For a broader coverage of the effect of natural resource exports on elements of the balance of payments, see Harding and Venables (2013).

[3] For example, Cashin et al. (2004) considers 19 commodity currencies during the period 1980-2002, and found estimates of long-run elasticities ranging between 0.16 for Iceland and 2.03 for Ecuador.

globalization? In addition, we look for differences in the short-run and long-run exchange rate responses to commodity price shocks.[4]

As a preview and to illustrate the large heterogeneity in the domestic currency responses to movements of world commodity prices, we consider a sample of 63 major commodity-exporting countries with significant commodity export shares (see Table A1). Regressing country-by-country the real effective exchange rate (*REER*) on real commodity price index (*RCP*) specific to each country's commodity export basket, we find the elasticity estimates to be significantly different from zero for 44 countries at the 5% level.[5] Figure 1 illustrates the range for these estimates. Overall, the *REER* and *RCP* are positively correlated for the majority of commodity-exporting countries in our sample with few exceptions. Elasticity estimates range from −2.93 (Argentina) to 4.79 (Ghana) with a median value of 1.28 (See column 6 in Table A1 in Appendix). In other words, when the world price of a country's commodity exports go up by 1%, some country's currency depreciates, while some others strengthens by over four-times in response.

INSERT FIGURE 1 HERE

What may account for this heterogeneity? We first note that since the 1980s, there have been important policy and structural changes in a large number of developing countries, as well as in the broader world commodity market environment. On the monetary policy front, for example, an increasing number of countries have adopted inflation targets, moved to greater exchange rate flexibility, and accumulated international reserves in the aftermath of a sequence of financial crises. Trade and financial integration progressed rapidly. In addition, a dramatic increase in commodity prices for the past decade, driven in large part by the strong global

[4] Chen et al. (2010) exploits short-run asset pricing dynamics, focusing on five developed economies including Australia, Canada, New Zealand, South Africa and Chile, although the majority of literature emphasize on the long-run relationship between commodity prices and the real exchange rates.

[5] The regressions include a time-trend and use Newey-West standard errors. In this paper, we use the real effective exchange rate (*REER*) as a measure of the international competitiveness of a country against all of its trade partners. This measure of real exchange rate provides an advantage to reduce potential biases associated with the choice of the anchor currency. An increase in the real effective exchange rate is defined as a real appreciation of the domestic currency relative to its trade partners. The real commodity price index (*RCP*) is defined as the world nominal price of country's commodity exports deflated by the price index of manufactured exports of industrial economies. More information about the *REER* and *RCP* including their construction and data sources are presented in Appendix. See Table A3 in Appendix for a list of commodities employed in the construction of *RCP*.

demand due to the rapid growth in emerging market economies and development of biofuels, affects the macroeconomic performance of energy and metal exporters more than food and raw materials exporters in general because the former typically depends more heavily on commodity exports and prices of industrial inputs are more sensitive to the global business cycle.

In order to understand the heterogeneous exchange rate responses to commodity prices across commodity exporters, we present the standard traded/nontraded goods model in the next section. Our model suggests that three factors affect the link between *REER* and *RCP*: the nation's export market structure (i.e., its degree of commodity export dependency and its export share in world markets), monetary policy choices (in the form of inflation-targeting, nominal exchange rate flexibility and international reserves management) and its degree of openness (both trade and financial). Our empirical results are largely consistent with the model predictions.

We also look for differences in commodity price elasticities over different time horizons. The majority of commodity currency literature focus on the long-run cointegrating relationship between the *REER* and *RCP* in a small set of developed countries.[6] Although standard theoretical channels such as the income effect of Dornbusch (1980) and the Balassa-Samuelson effect (Balassa, 1964; Samuelson, 1964) suggest the link between these two relative prices may hold only in the long-run, a short-run relationship may exist.[7] Chen and Rogoff (2003) report significant and positive short-run commodity price elasticities in Australia and New Zealand. We test if we can observe any differences in the commodity currency phenomenon between in the short-run and in the long-run using a panel data set of sixty three commodity exporters at different stages of development and in various economic environments.

Our non-stationary panel data estimation results demonstrate a strong long-run cointegrating relationship between the *REER* and *RCP*, in line with the "commodity currency" country experience in the literature. These results are robust to the structural shift consideration.

[6] There are a few recent papers whose analysis is based on a panel of developed and developing commodity exporters. These include Ricci et al. (2008), Coudert et al. (2008), and Bodart et al. (2011, 2012).

[7] For instance, consider a commodity exporter that experiences a resource-based export boom. This typically leads to a balance-of-payment surplus and an accumulation of international reserves. When monetary authorities attempt to sterilize an increase in money supply, it could lead to an increase in the monetary base and national price level if imperfectly done. It is possible in the short run to have inflation rate that is higher than required to achieve the real appreciation generated by export boom. This is one of the cases that could occur in the short-run as a real appreciation episode in response to a rise in commodity prices (Edwards, 1986).

However, in contrast to previous studies based on the currencies of Australia and New Zealand, we find weak evidence of the commodity-currency phenomenon in the short-run in our data. Our results also reveal that the magnitude of *REER* response depends on the export structure and policy choices of a country. In particular, the reaction of the real exchange rate to a commodity boom would be amplified if a country is characterized by any of the following traits: i) heavy commodity export dependency, ii) monopoly pricing power by possessing a dominant export share in world markets, iii) fixed nominal exchange rate, iv) inflation-targeting, v) low international reserves holdings, vi) open financial market and vii) low degree of trade dependence.

While high commodity prices of any type bring about higher export revenue for a country that exports a particular commodity, it may also lead to the inflationary pressure, inflow of large hot money, and deterioration of the price competitiveness of non-commodity sectors in the world trade market. Therefore, effectively managing these adverse consequences of commodity price fluctuations is a natural interest of policy makers in commodity exporting countries. We believe that findings in this paper help them to find appropriate policy responses to stabilize their economy by effectively dampening rather than amplifying the commodity price shocks.

The remainder of this paper is organized as follows. Section 2 sets out the structural model that examines the theoretical factors influencing the commodity price elasticity of real exchange rate. Section 3 explains the estimation procedure including an empirical model specification and data diagnosis. Section 4 presents the long-run and short-run estimation results based on our non-stationary panel data set and robustness of the results. Section 5 concludes.

2 Determinants of commodity price elasticities

In this section, we present a theoretical model that highlights the impact of commodity terms of trade on the real exchange rate in a commodity exporting country.[8] The model allows us

[8] For the purpose of our work, we adopt the models presented in Obstfeld and Rogoff (1996) and Cashin et al. (2004).

to discuss potential factors that may affect real exchange rate responses to global commodity price shocks.

2.1 The model

2.1.1 Production

Consider a small open economy that produces two types of goods, exportables "*X*" and nontradables "*N*". In our framework, exportables "*X*" are primary commodities unless otherwise noted. We assume primary commodities are all exported. With constant-returns to scale production technology, competitive domestic firms produce two types of goods using labor (L) and capital (K):

$$Y_N = A_N L_N^{\alpha} K_N^{1-\alpha} \tag{1}$$

$$Y_X = A_X L_X^{\beta} K_X^{1-\beta} \tag{2}$$

where Y_N and Y_X are output of the nontradable and exportable goods; $A_i, L_i,$ and K_i are productivity shocks, labor and capital in sector i, $i = N, X$; $0 < \alpha < 1$ and $0 < \beta < 1$. Capital is allowed to move between sectors and countries, but labor is assumed to be mobile only between sectors within the economy. The total domestic labor supply is inelastically given by $L = L_N + L_X$. Because capital is internationally mobile, the domestic marginal product of capital is given by the world interest rate (r^*), while perfect labor mobility between industries ensures wage (w) equalization across sectors.

Let p_X be the world price of exportables exogeneously given to the small economy, and p_N be the domestic price of non-traded goods. We assume that the law of one price holds for the exportable goods so that:

$$Ep_X = p_X^* \tag{3}$$

where E is the nominal exchange rate, defined as the price of domestic currency in terms of foreign currency, and an asterisk denotes a foreign value. Let us also define the capital-labor ratio in both sectors as $k_i \equiv K_i/L_i$, i = N, X. Then, the firm's profit maximizing first-order conditions for labor and capital in both sectors are given by the following functions:

$$\alpha p_N A_N k_N^{1-\alpha} = w \quad (4a)$$
$$(1-\alpha) p_N A_N k_N^{-\alpha} = r \quad (4b)$$

in the nontradable sector and

$$\beta p_X A_X k_X^{1-\beta} = w \quad (5a)$$
$$(1-\beta) p_X A_X k_X^{-\beta} = r \quad (5b)$$

in the exportable sector. By combining first-order conditions in each sector, we derive zero-profit conditions for both sectors:

$$p_N A_N k_N^{1-\alpha} = w + r k_N \quad (6)$$
$$p_X A_X k_X^{1-\beta} = w + r k_X \quad (7)$$

Taking a log-differentiation of equations (6) and (7) making use of (4b) and (5b) yields

$$\hat{p}_N = \frac{\mu_{LN}}{\mu_{LX}}\left(\hat{p}_X + \hat{A}_X\right) - \hat{A}_N \quad (8)$$

where a hat above the variable denotes a logarithmic derivative and $\mu_{LN} \equiv wL_N/p_N Y_N$ and $\mu_{LX} \equiv wL_X/p_X Y_X$ be labor income share in the nontraded and exportable sectors, respectively. Empirically, nontraded goods tend to be at least as labor-intensive as exportable goods. Thus, it is standard to assume $\frac{\mu_{LN}}{\mu_{LX}} \geq 1.$ Equation (8) states that, along the perfect foresight path, a positive (commodity) terms-of-trade shock will lead to an increase in relative price of nontradables and this increase will be larger the more labor-intensive are nontradables relative to exportables. The equation (8) also reflects the Harrod-Balassa-Samuelson effect by showing a positive relationship between a productivity shock in the exportable sector and the price of nontradable sector under the assumption of free labor mobility. With perfect international capital mobility and free movements of labor and capital across sectors, the relative price of nontradables is entirely determined by the production side of the model and is independent of demand side factors.

2.1.2 Consumption

The economy is inhabited by a continuum of identical individuals that supply labor and consume two goods – nontradables (*N*) and importables (*T*). A representative consumer's utility function takes the Cobb-Douglas form:

$$U = \tau C_N^{\gamma} C_T^{1-\gamma} \tag{9}$$

where C_i is consumption of good i, i = N, T; $0 < \gamma < 1$ and $\tau = 1/\gamma^{\gamma}(1-\gamma)^{(1-\gamma)}$. Hence, consumption-based consumer price index (p) is given by

$$p = p_N^{\gamma} p_T^{1-\gamma} \tag{10}$$

2.1.3 How is the real exchange rate determined?

We now define the real exchange rate as the foreign price of the domestic basket of consumption (Ep) relative to the foreign consumer price index in foreign currency (p^*).[9] Using (10), we can write the real exchange rate (q) as

$$q = \frac{Ep}{p^*} = \frac{E p_N^{\gamma} p_T^{1-\gamma}}{\left(p_N^*\right)^{\gamma^*} \left(p_T^*\right)^{1-\gamma^*}} = \left(E \frac{p_N}{p_N^*}\right)^{\gamma} \tag{11}$$

[9] Following Cashin et al. (2004), we assume that the foreign economy produces nontraded (*N*), intermediate (*I*), and final traded (*T*) goods using labor and capital. The foreign country is a manufacturing good exporter using the commodity as an input that is imported from the domestic economy. From the production side, it is straightforward to show that the price of foreign non-traded goods as a function of foreign intermediate goods. If we assume that the foreign final goods sector uses v share of intermediate goods and (1 - v) share of commodities imported from the domestic economy to produce the final traded goods, the cost minimization leads to the following per unit cost: $p_T^* = (p_I^*)^{v}(p_X^*)^{1-v}$. Thus, the implied CPI is: $p^* = (p_N^*)^{\gamma^*}(p_T^*)^{1-\gamma^*}$. We also assume that foreign economy's share of commodity import is small (i.e., v is a relatively large fraction) enough that world price changes in commodities have negligible effect on the foreign currency value.

assuming that $\gamma = \gamma^*$ and the law of one price for importables $Ep_T = p_T^*$. Taking a log-differentiation of equation (11) gives

$$\hat{q} = \gamma\left(\hat{E} + f\left(\hat{p}_X\right) - \hat{p}_N^*\right) \tag{12}$$

when $\hat{p}_N$ is written as a linear function of $\hat{p}_X$ using (8). The equation (12) shows that, given $\hat{p}_N^*$, the real exchange rate in the home country appreciates in response to an increase in the price of exportables, with the extent of this appreciation depending on the variables/parameters present in the equation.

2.2 Factors influencing the commodity price elasticity

2.2.1 Degree of openness

Trade Openness (*TO*). Since $\hat{p}_N$ and $\hat{p}_X$ move in a proportionate fashion with a given relative labor income share as shown in (8), the equation (12) shows that the elasticity of the real exchange rate with respect to the price of exportables depends on γ, where γ captures a share of nontradables in a basket of domestic consumption. Therefore, if the economy depends heavily on imported goods with large $(1-\gamma)$, the real exchange rate response to an increase in the price of exportable commodities would be relatively small. This makes sense intuitively because the domestic price of importables is likely to be higher if a country facing a positive commodity price shock places greater limits on the volume of its imports, leading to the higher overall domestic price level and the more appreciation pressure on the real exchange rate.

Financial Openness (*FO*). One of the critical assumptions made in our theoretical model is to allow capital to be mobile internationally and between sectors. With the exogenously given return to capital (and price of exportables), wages are entirely determined by (7). This wage in turn determines the domestic price of nontradables according to (6). A commodity price boom attracts foreign capital to the exportable goods sector, raising marginal productivity of labor and wages of that sector by (5a). Now through the channels described in (6) and (7), this should eventually generate an increase in price of nontradables according to (8), resulting in an equilibrium real exchange rate appreciation. Suppose now that foreign capital inflow is restricted in a commodity-exporting country and, as a result, supply of capital is not perfectly elastic any longer. Real appreciation pressure caused by a positive price shock in exportables would be relatively small under a capital control as improvement of the marginal productivity of labor in exportables is likely to be constrained with inelastic supply of capital. This in turn leads to a relatively smaller increase in wages and price of nontradables, putting relatively smaller appreciation pressure on the value of domestic currencies. Therefore, a positive relationship may exist between the degree of financial openness and commodity price elasticity.

2.2.2 Export market structure

Commodity Export Dependency (*CEX*). Let us introduce the new sector called the exportable manufacturing sector to our domestic production side of the model. Now the domestic production is given by:

$$Y = Y_N^{\rho} Y_{XC}^{\delta} Y_{XM}^{1-\rho-\delta} \tag{13}$$

where subscripts *XC* and *XM* denote exportable commodity and exportable manufacturing sectors, respectively. It is straightforward to show that the price of the one unit of composite domestic product is given by:

$$p = p_N^{\rho} p_{XC}^{\delta} p_{XM}^{1-\rho-\delta} \tag{14}$$

Equation (14) indicates that when a share of exportable commodity δ in the domestic production is large, the real exchange rate response to a commodity terms-of-trade shock that is exogenously given to the domestic economy is likely to be amplified.

World Market Share (*MSH*). In our model above, we assume that the domestic economy is so small that it takes the price of exporting commodities from the rest of the world. This assumption may not hold if a country has a dominant world market share of commodity products and, as a result, has some degree of market power.[10] Consider a country that has monopoly

[10] Examples of countries with a sufficiently large market share of commodity exports include Chile (copper), Malaysia (palm oil), Philippines (coconut oil), and Saudi Arabia (crude oil). Each of these countries often accounts

power in the world market for a certain commodity in the sense that a large volume of exports places downward pressure on the world price. Domestic producers of such country have an incentive to expand production and export more to increase their revenue if a domestic currency depreciates. The world price of the commodity is expected to fall due to the large supply. Based on this logic, we can derive the production of exportables as a negative function of real exchange rate:

$$Y_X = Y_X\left(\overset{(-)}{\frac{Ep}{p^*}}\right) \tag{15}$$

Furthermore, in the export market of this commodity, the world price depends on the supply of the world leading producer. Therefore, the world commodity price is given by:

$$p_X^* = p_X^*\left(\overset{(-)}{Y_X}\right) \tag{16}$$

holding everything else that possibly influences the price of exportables constant. Thus, from equations (12) and (15), we find that a commodity boom appreciates the country's currency, and as this squeezes its exports, the supply of commodity in the world export market falls. But the reduction of exports pushes up the world price further by (16) as, by assumption, the country is large. Consequently, the appreciation pressure under a commodity boom can be amplified when the commodity prices are endogenously determined.

2.2.3 Monetary policy choices.

Monetary policy options that may drive the movements of the nominal exchange rate and domestic price level are also important candidates affecting the commodity price elasticity as E and p_N appear in equation (11). With our assumption of perfect cross-border capital mobility, a monetary authority faces a trade-off between pegging the exchange rate and keeping autonomous monetary policy decisions. This is due to the well-known open-economy trilemma; the three desirable macroeconomic objectives, namely, exchange rate stabilization, capital mobility, and

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for more than one third of world production of its primary commodity.

domestic monetary autonomy cannot be jointly pursued.[11] Instead, only two out of three objectives can be compatible. Since the early 1990s, a number of countries have adopted an inflation targeting scheme and an exchange rate policy that allows a greater flexibility. Furthermore, a series of financial crises around the globe in the 1990s have triggered many developing countries, crisis inflicted countries in particular, to accumulate a sizeable stock of international reserves to prevent the recurrence of financial turmoil by preserving extra liquidity. In what follows, we discuss about the possible effect of monetary/exchange rate policy choices on the real exchange rate movements in commodity exporting countries.

Inflation Targeting (*IT*). A central bank in *IT* countries sets a target inflation rate, which is defined by a mean value or a range along with a time horizon, and adjusts the nominal interest rate to achieve its policy goal that includes not only price stability but also output and employment stability. However, as implied by the macroeconomic trilemma, the central bank cannot simultaneously achieve monetary independence and a hard peg. This means that, in theory, a country with an inflation targeting regime should allow a considerable degree of exchange rate flexibility.[12] As a result, it is a challenging task to determine if the commodity price elasticity of the real exchange rate is bigger under an *IT* regime or non-*IT* regime, because we do not know a priori how much the nominal exchange rate should adjust under the commodity price boom when the *IT* regime is in effect.[13] Furthermore, as long as exchange rate movements drive inflation fluctuations, the central bank values exchange rates as an important element in an *IT* regime when making monetary policy decisions. Empirical findings in Edwards (2006) support this view and find no evidence that the adoption of *IT* increases real effective exchange rate volatility in Australia, Canada, Chile and Israel. Instead, he finds opposite episodes in Chile and Israel: real exchange rate volatility tends to be lower in these inflation targetors during the period 1988-2005. In summary, the question of whether the response of the

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[11] See Shambaugh (2004) and Obstfeld et al. (2005) that discuss about the presence of open-economy trilemma across different regimes and countries.

[12] This is discussed by many authors in the literature (See, for example, Masson, Savastano and Sharma, 1997; Mishkin and Savastano, 2001; and Brenner and Sokoler, 2010).

[13] Although our model emphasizes, for a simplicity purpose, a commodity price transmission to the real exchange rate through adjustments in price of nontradables, a nominal exchange rate adjustment channel should not be overlooked. In fact, Chen (2002) finds commodity currencies in Australia, Canada, and New Zealand using their nominal exchange rates.

real exchange rate is weaker in the *IT*-countries than in non-*IT* countries should be answered based on the empirical findings.

Exchange Rate Regime (*EXR*). The Mundell-Fleming model predicts that under a credible fixed exchange regime and free capital mobility, the central bank loses an ability to make autonomous adjustments in monetary policy: the (risk-adjusted) domestic interest rate must be equal to the foreign interest rate. If a commodity price transmission into the real exchange rate works mostly through nominal exchange rate adjustments with a limited role played by domestic price changes, pegging exchange rates would be an appropriate policy instrument to stabilize the real exchange rate. If the opposite is true, however, a commodity boom may lead to a greater real appreciation pressure under the peg. Empirical evidence in the literature is mixed. Broda (2004) shows that in response to a decline in terms-of-trade, the real exchange rate depreciation is small and slow in pegs but large and immediate in floats. Aizenman et al. (2012) also find that the effect of commodity terms-of-trade shocks on the real effective exchange rates is much larger under the floating regime. On the contrary, Bodart et al. (2011) shows that the elasticity of real effective exchange rate with respect to the price of main commodity exports is lower in floats. Therefore, our theoretical judgement of the effect of a peg on the real exchange rate in response to a commodity price shock is uncertain.

International Reserves (*RES*). Over the last two decades, an increasing number of developing countries in the world have chosen foreign exchange interventions through massive international reserve accumulation. Switching to a regime that allows a greater exchange flexibility during the same period has not stopped this reserve stockpiling behavior. Indeed, reserves to GDP ratios have dramatically increased especially in the aftermath of East Asian financial crises. While a major motive in the literature about the large international reserves hoarding is to reduce the probability of a sudden stop of capital flows (Aizenman and Marion, 2003; Jeanne and Ranciere, 2011), it may also help lower the transmission of commodity price shocks into real exchange rates. For example, commodity-dependent Latin American countries appear to effectively lower volatility of the real effective exchange rates in the face of commodity terms-of-trade shocks using ample reserves (Aizenman et al., 2012). It is our interest to test if large international reserve holdings help to stabilize the real exchange rate in commodity exporting countries.

The theoretical effect of factors influencing the commodity price elasticity of the real exchange rate $\left(\frac{\hat{q}}{\hat{p}_X}\right)$ discussed in this section can be summarized as follows:

$$\frac{\hat{q}}{\hat{p}_X} = f\left(\overset{-}{TO}, \overset{+}{FO}, \overset{+}{CEX}, \overset{+}{MSH}, \overset{+/-}{IT}, \overset{+/-}{EXR}, \overset{-}{RES}\right) \tag{17}$$

where *TO*, *FO*, *CEX*, *MSH*, *IT*, *EXR* and *RES* denote trade openness, financial openness, commodity export dependency, world market share, inflation targeting, exchange rate regime and international reserves, respectively. The signs above the variables indicate the expected effect of these variables on the commodity price elasticity.

3 Empirical implementation

3.1 Model setup

To empirically test the above theoretical determinants of commodity price elasticities of real exchange rate, we begin with the standard regression model used in the commodity currency literature:

$$REER_t = \alpha_1 + \alpha_2 RCP_t + \varepsilon_t \tag{18}$$

where $REER_t$ and RCP_t are the real effective exchange rate and real commodity price index respectively for each country and the error term ε_t is i.i.d. over periods. The parameter that determines whether a country has a commodity currency is α_2. Our goal is to identify potential factors that may influence a relationship between the *REER* and *RCP*, which is measured by α_2. More formally, we want to explain the parameter α_2 using a set of variables *X* such that

$$\hat{\alpha}_{2i} = g(X_i) = \beta_1 + \boldsymbol{X}_i \boldsymbol{\beta}_2 \tag{19}$$

where $\boldsymbol{X}$ includes seven factors of commodity currencies introduced in the previous section and $\boldsymbol{\beta}_2$ is a vector of coefficients. Combining (18) and (19), under the exogeneity assumption, the model for our empirical estimation is[14]:

$$REER_{it} = \alpha_1 + \beta_1 RCP_{it} + RCP_{it}\boldsymbol{X}_{it}\boldsymbol{\beta}_2 + u_{it} \tag{20}$$

From the model (20), we know that β_1 is the elasticity of the *REER* with respect to *RCP* and $\boldsymbol{\beta}_2$ measures marginal impact of *RCP* changes on *REER* conditional on structural/policy factors *X*.[15] Our primary interest centers on the coefficient vector $\boldsymbol{\beta}_2$: a significant positive coefficient implies that a given positive *RCP* shock puts larger appreciation pressure on the *REER* conditional on structural/policy factors.

3.2 Variable Definitions

3.2.1 Market structure measures

Commodity Export Dependency (*CEX*). We define a country's commodity export dependency as follows: for each country *i* and time *t*, CEX_{it} = total commodity exports$_{it}$ / total merchandise exports$_{it}$. A high value of *CEX* indicates a country's heavy reliance on commodity exports and a low degree of export diversification.

World Market Share (*MSH*). We introduce a world market share of commodity exports as a proxy for market power. So for each country *i*, commodity *j*, and time *t*, the world market share is defined as $MSH_{it} = \sum_{j=1}^{J} W_{ijt}(\text{market share}_{ijt})$, where W_{ijt} = commodity exports$_{ijt}$ / total commodity exports$_{it}$ and *market share*$_{ijt}$ = commodity exports$_{ijt}$ / world supply of commodity$_{jt}$. Since a country's export basket typically includes multiple commodities, we construct a weighted index of market share to better identify the impact of a country's potential pricing power on its corresponding (aggregate) commodity prices.

3.2.2 Monetary policy measures

[14] The model (18) is in a time-series dimension while the model (19) in a cross-sectional dimension. We admit that combining these two models into a single panel model is unjustifiable. However, presenting these intermediate steps helps to understand our underlying idea about the estimation strategy.

[15] Note that *REER* and *RCP* are in logarithm in all of our empirical procedures.

Inflation Targeting (*IT*). Since the 1990s, a number of central banks in both developed and developing economies have adopted inflation targeting (*IT*) as an instrument to achieve the low and stable average inflation. Commodity exporters were not an exception.[16] The exact adoption dates of inflation targeting are from Roger (2009).

Exchange Rate Regime (*EXR*). In order to study the effect of nominal exchange rate flexibility, we follow Ilzetzki et al. (2008; called IRR hereafter) and use their coarse classification for a country's exchange rate regime choices.[17] This has six regimes, namely, hard peg, soft peg, managed floating, freely floating, freely falling, and dual market.[18] The larger the code, the more flexible the regime is. Countries with the hard and soft pegs (IRR code = 1 and 2) are defined as fixed exchange rate regime economies.

International Reserves (*RES*). We extract data for international reserves from IMF IFS and nominal GDP from World Bank WDI to construct a *RES* (= international reserves / NGDP) variable.

3.2.3 Openness measures

Financial Openness (*FO*). Financial openness represents a country's degree of capital account openness. In order to measure a country's degree of capital account openness, we use Chinn-Ito index (2008).[19] This index measures "the extent and intensity of capital controls based on the information from the IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions* (AREAER)."[20] The index runs from -1.84 to 2.48, where higher values indicate that a country is more open to cross-border capital transactions.

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[16] See Appendix Table A4 for inflation targeting adopting countries and adoption dates in our sample.
[17] An updated classification is obtained from Ilzetzki's webpage at http://personal.lse.ac.uk/ilzetzki/IRRBack.htm.
[18] See Table A5 in Appendix for details.
[19] A data set for financial openness index is from http://web.pdx.edu/~ito/Chinn-Ito_website.htm.
[20] Published annually since 1967, the *AREAER* offers a summary table with binary indicators for four types of de facto controls: (i) multiple exchange rates, (ii) restrictions on current account transactions, (iii) restrictions on capital account transactions, and (iv) regulatory requirements of the surrender of export proceeds. In 1998, the *AREAER* expanded the four subcategories and now offers fourteen binary indicators for de facto controls on: capital market securities, collective investment instruments, commercial credits, foreign direct investment, and real estate transactions among others. Chinn-Ito index (2008) is an intensity-modified index of capital controls by taking all four types of controls into account instead of focusing only on capital account transaction controls.

Trade Openness (*TO*). Trade openness measures the degree of trade dependency reflecting how much the economy relies on tradable goods. We use the ratio of exports plus imports to GDP as a measure of trade dependence in our empirical procedure.[21]

3.3 Data description and characteristics

Our empirical analysis is based on a quarterly panel data set of 63 commodity exporters during the period from 1980q1 to 2010q4. World prices and the individual country-level export volume of 58 commodities are used in constructing real commodity prices and other relevant commodity export structural factors. Description of variables and their sources are presented in Table A2 in Appendix. A majority of our control variables are available only at an annual frequency and interpolated to a quarterly frequency through "constant-match average". For our main variables *REER* and *RCP*, which are available at a monthly frequency, we use "last observation" method that sets the quarterly observation equal to the value in the last of the corresponding monthly observations. Table A6 in Appendix contains descriptive statistics for the sample data used in our empirical estimation.

As an initial data diagnosis, we show in figure 2 the time-series properties of the *REER* and *RCP* using a small set of countries from our data. Six countries from Figure 1 (and from Table A1 in Appendix) are selected based on the size of the exchange rate response to commodity price fluctuations. Ghana and India are the countries with the highest elasticity estimates, the Canada and Burundi have median elasticity estimates, and Argentina and Peru have the lowest elasticity estimates.

Visual inspection of the figure suggests that each of the *REER* and *RCP* does not appear to move around a given long-run equilibrium level, suggesting the potential presence of unit-roots in both series. Despite wild fluctuations of the exchange rate and commodity prices individually, we observe a close co-movement between these two series over a long period of time for the majority of the commodity exporters in our sample, except for a few cases such as Argentina and Peru. Furthermore, the relationship between the *REER* and *RCP* exhibits structural

[21] In the theoretical model in section 2, we use a share of importables $(1-\gamma)$ in a representative consumer's consumption bundle to capture the degree of trade openness. Our empirical estimation results are robust to the use of this theory-consistent definition of trade openness (namely, import-to-NGDP ratio) and are available upon request.

shifts in some countries including Ghana and Peru. For example, the *REER* of Ghana experienced steep depreciation from the period after the Structural Adjustment Programme (SAP) in 1983, which included exchange rate reforms until 1990. On the other hand, Peru experienced dramatic appreciation of its domestic currency because of the hyperinflation episodes in the late 1980's. We therefore take a country's potential regime shift into account as a robustness check.

INSERT FIGURE 2 HERE

In addition to these time series properties of data, cross-sectional dependence is likely to be important in our data because common shocks such as global recession and spillover effects could affect the *REER* of trade partners as a group. Moreover, by the nature of its construction, *REER*s are interdependent between trade partners.

3.3.1 Order of integration: Panel unit-root tests

To formally test whether the *REER* and *RCP* are stationary or not, we consider three panel unit root tests: the Levin-Lin-Chu (2002) (LLC) test, Im-Pesaran-Shin (2003) (IPS) test, and Pesaran's (2007) cross-sectionally augmented IPS (CIPS) test.[22] The first two are the popular first generation panel unit root tests based on the ADF (augmented Dicky-Fuller) test. Formally, their base-line regression model takes the form,

$$\Delta y_{it} = \alpha_i + \theta_t + \beta_i t + \rho_i y_{it-1} + \varepsilon_{it} \tag{21}$$

where the errors ε_{it} (i = 1, 2, …, N, t = 1, 2, …, T) are independently distributed across both i and t, with zero means and finite heterogeneous variances, σ_i^2.

The LLC test assumes that all cross-sectional units have a common autoregressive parameter although it allows for individual effects, time effects and a time trend. The test may be viewed a pooled Dickey-Fuller (DF) or Augmented Dickey-Fuller (ADF) test with the null hypothesis of $H_0: \rho_i = 0 \forall_i$, against the alternative that all series are stationary, $H_1: \rho_i < 0 \forall_i$.

[22] Theoretical background for each of panel non-stationary and cointegration test is from Baltagi (1995). See Chapter 12 of his book and references therein for more details.

This restrictive assumption is relaxed in the IPS test by allowing only a fraction (but not all) of the panel is stationary under the alternative (H_1: $\rho_i < 0$ for at least one i). However, the first generation tests ignore the cross-sectional dependence amongst units included in the sample, possibly generating significant size distortion in test statistics.

Cross-sectional dependence is likely to be important in our case as explained earlier. Driscoll and Kraay (1998) note that standard error estimates of commonly applied co-variance matrix estimation techniques such as OLS, White, and Rogers by erroneously ignoring spatial correlation in panel regressions are biased and hence statistical inference that is based on such standard error is invalid. Typically, ignoring cross-sectional dependence leads to overly optimistic standard error estimates. Furthermore, O'Connell (1998) highlights the importance of accounting for cross-sectional dependence when performing a unit-root test in panels of real exchange rates. Thus, we conduct the cross-sectional dependence (CD) test by Pesaran (2004). Test results in table 1 reject the null of cross-sectional independence at the 1% significance level, indicating that the regression residuals are cross-sectional dependent.

INSERT TABLE 1 HERE

This finding motivates the implementation of the second generation panel unit-root test allowing for cross-sectional dependence (Pesaran, 2007). To eliminate the cross-sectional dependence, the ADF regressions are augmented with the cross-section average of lagged levels and first-differences of individual series. The CIPS test is based on the following cross-sectionally augmented Dicky-Fuller regression:

$$\Delta y_{it} = \alpha_i + \theta_t + \beta_i t + \rho_i y_{it-1} + \gamma_0 \bar{y}_{t-1} + \gamma_1 \Delta \bar{y}_t + \varepsilon_{it} \tag{22}$$

where $\bar{y}_{t-1}$ and $\Delta \bar{y}_t$ are proxies for the unobserved common factors. Parallel to the IPS test, heterogeneity in the autoregressive coefficient is allowed under the alternative hypothesis (H_1: $\rho_i < 0$ for at least one i) and the test is based on the mean of individual DF (or ADF) t-statistics of each unit in the panel. All tests are normally distributed under the null hypothesis of non-stationarity. Panel unit-root test results are reported in Table 2, informing that we strongly reject the null of unit-root and both *REER* and *RCP* series are integrated of order one.

INSERT TABLE 2 HERE

3.3.2 Panel cointegration tests

If two variables are found to be integrated of order greater or equal to one, then it could be the case that the two series are cointegrated. Figure 2 also indicates the possible (long-run) common trend between *REER* and *RCP*. In this subsection, we implement three panel cointegration tests to see whether there indeed exists the common trend. We apply Kao (1999), Pedroni (2004), and Westerlund (2007) panel conintegration tests. Under the null hypothesis of no cointegration, the Kao (1999) test assumes cointegrating vectors to be homogeneous across individuals while Pedroni (2004) allows heterogeneity in the cointegrating vectors as well as fixed effects and trends in the data. Among the seven Pedroni's tests, four are based on within-dimension and the rest are based on between-dimension. Alternative hypotheses for Pedroni's are common AR coefficients for within-dimension and individual AR coefficients for between-dimension. The two cointegration tests mentioned above do not allow for the presence of cross-sectional correlation that is detected by the Pesaran (2004)'s CD test. Thus, we proceed with the Westerlund (2007) error-correction-based cointegration test. It is designed to test for the absence of cointegration by determining whether there exists error correction for individual countries or for the panel as a whole. This test is more general than Pedroni (2004) by allowing for a large degree of heterogeneity and dependence within as well as across the cross-sectional units. As shown in table 3, the overall result is in favor of cointegration, suggesting existence of long-run relation between *REER* and *RCP*.

INSERT TABLE 3 HERE

3.4 Additional controls

We estimate equation (19) using the (country) fixed effect model to correct for omitted variable bias caused by unobserved heterogeneity. In addition, fixed effects are necessary in our case because *REER* measures are country-specific indexes, making a cross-country comparison impossible. Furthermore, in order to avoid a potential identification problem resulting from ignoring cross-sectional dependence, we report Driscoll and Kraay (1998) standard errors to correct for spatial correlation, autocorrelation and heteroskedasticity throughout our estimation procedures.[23] Conditional factors are converted into binary dummy variables based on the threshold of sample median for each series.[24]

3.5 Long-term vs. short-term

3.5.1 Long-run estimation method: Dynamic OLS

Recognizing non-stationarity and the presence of cointegration for *REER* and *RCP*, we apply the panel DOLS (Dynamic Ordinary Least Squares) with 1 lead and 1 lag to estimate the cointegrating parameter. For country *i* and for time period *t*, the long-run estimation is done based on the model

$$REER_{it} = \alpha_1 + \beta_1 RCP_{it} + \sum_{s=-p}^{p} \delta_s \Delta RCP_{it+s} + RCP_{it}\boldsymbol{X}_{it}\boldsymbol{\beta}_2 + \theta a_i + u_{it} \tag{23}$$

where β_1 is the long-run cointegrating coefficient, Δ denotes the first-difference operator, δ_s is a coefficient vector of leads and lags of the changes in real commodity price index, X_{it} is a set of structural/policy factors, a_i is the country fixed effect, and u_{it} is the disturbance term. Country fixed effects are included in the model to capture the unobserved country-specific factors, reducing the omitted variable bias.

3.5.2 Short-run estimation method: First difference

We employ a simple first difference model to study the short-run relationship between *REER* and *RCP*. The model takes the following form:

$$\Delta REER_{it} = \beta_0 + \beta_1 \Delta RCP_{it} + (\Delta RCP_{it})\boldsymbol{X_{it}\beta_2} + \varepsilon_{it} \tag{24}$$

---

[23] Although it is not presented in this paper, we found the substantial cross-sectional dependence even in the first-differenced real effective exchange rate regression model.

[24] For example, we set *CEX* = 1 if *CEX* > median(*CEX*) and *CEX* = 0 otherwise. For a trade dependency measure, we use the threshold set in Aizenman et al. (2012): highly trade dependent when a ratio of trade (= *EX* + *IM*) to 2×*NGDP* is greater than 0.3.

where X_{it} is a vector of seven control factors.

4 What makes a commodity currency?

4.1 Commodity currency in the long-run

First of all, we find the strong and robust link between *REER* and *RCP* in the long-run across the different specifications from (1) to (5) in table 4. Indeed, there is almost one-for-one cointegrating relation between them and this relation is strongly significant. For example, one percent permanent increase in commodity price index will cause the real effective exchange rate to appreciate by 0.929 percent according to the specification (1). The effect of globalization is presented in column (2). Higher trade dependence tends to dampen the effect of *RCP* shocks on *REER* while a greater degree of financial openness is found to be amplifying the shock, in accordance with the theory. For policy variables in column (3), we find the response of *REER* to the commodity price fluctuations tends to be larger under a peg and smaller under large international reserve holdings. This suggests that the size of commodity terms-of-trade pass-through onto domestic price level appears to dominate the nominal exchange rate stabilization under a peg. Our results also confirm the buffering role of foreign reserves in mitigating the impact of external shocks on the real exchange rate. Results in column (4) show that the larger the share of the commodity exports in total exports, the larger the commodity price elasticity of the real exchange rate, consistent with the empirical finding by Bodart et al. (2012). From the market share interaction variable, the positive coefficient indicates that we find some evidence of monopoly pricing power in our data. Estimation results including all conditional variables are reported in column (5) as a robustness check.

INSERT TABLE 4 HERE

The interaction variable model increases the likelihood of the presence of multicollinearity issues. In the presence of a high multicollinearity, it is often the case that the regressors of primary interest are jointly uninformative. We thus perform a *F*-test for each specification to see if it is a valid procedure to add interaction terms into our regression model.

So, for example, the null hypothesis of a F-test for the specification (2) in Table 4 is $H_0: \beta_{RCP*TO} = \beta_{RCP*FO} = 0$ and the alternative is H_1: at least one $\beta \neq 0$. For the all specifications in the DOLS(1,1) estimation in Table 4, we find that the data overwhelmingly reject the joint null hypothesis and conclude that interaction terms are jointly significant and informative in explaining the long-run *REER* behavior.

4.2 Commodity currency in the short-run

Table 5 shows the first difference estimation results. In contrast to the literature (Chen and Rogoff, 2003), we find relatively weak evidence of short-run relationship between the real exchange rate and commodity prices in our full sample with the poor goodness-of-fit measures. One plausible explanation for this weak result is that there is a greater likelihood of foreign exchange intervention in developing countries that are more vulnerable to external shocks. Furthermore, all factors are indeed sector specific in the short-run, making the factor price adjustment nearly impossible.

Nevertheless, one variable that is worth paying attention here is a choice of inflation targeting. Countries with a narrow band of inflation rate require a relatively higher nominal exchange rate adjustment in response to the commodity price shock and this is reflected in the positive coefficient of the *IT* interaction term.

INSERT TABLE 5 HERE

4.3 Robustness checks

In this subsection, we want to check the robustness of our main results. We focus only on the robustness of our long-run estimation results.

First, we consider the potential presence of structural shifts in the long-run cointegrating relationship between *REER* and *RCP*. During the sample period between 1980 and 2010, developing countries in the globe experienced a significant structural change in policies including nominal exchange rate regime and inflation control, and wild fluctuations of world prices of commodities. All of these have a potential to affect the level of the real exchange rate and consequently a cointegrating relationship between *REER* and *RCP*. Abrupt changes in real effective exchange rate would blur the cointegrating relationship between *REER* and *RCP* and needs to be properly controlled. Gregory-Hansen (1996) propose a cointegration test that allows for regime shifts at an unknown point in time. We consider a level shift in the long-run relationship between *REER* and *RCP* as government interventions in developing countries typically aim at affecting the level of the real exchange rate. Gregory-Hansen (1996) test is based on the following model:

$$REER_t = \beta_0 + \beta_1 \vartheta_t + \beta_2 RCP_t + \varepsilon_t \tag{25}$$

where β_0 is the intercept in the original cointegrating relationship, β_1 is coefficient of the dummy variable ϑ_t that models structural change as follows:

$$\vartheta_t = \begin{cases} 0 & if \quad t \leq [T\pi] \\ 1 & if \quad t > [T\pi] \end{cases} \tag{26}$$

where *T* is the sample size, $\pi \in (0,1)$ is a fraction parameter that determines a timing of the level shift and [ ] denotes an integer part. Following Gregory and Hansen (1996), we use trim (π) of 0.15, which specifies the fraction of the data range that skips either end when examining possible break points. The test is applied to each country to detect shift dates, allowing for a level shift and lag length chosen based on AIC. The null hypothesis of the test is no cointegration against the alternative of cointegration with a single shift at an unknown point in time. The last two columns of Table A1 in Appendix report the Gregory-Hansen *Z*(t) statistics and the shift dates selected by the test. Selected shift dates are largely associated with the country specific macroeconomic events such as hyperinflation, exchange rate crisis, nominal exchange rate adjustment program including remarkable devaluation of the CFA franc by 50 percent in early 1994. The regression results presented in Table 6 include country specific dummy variables controlling for the structural shift dates. After controlling for regime shifts in the cointegrating relationship, the DOLS(1,1) regression results reported in Table 6 are still very similar to the main results presented in Table 4. Thus, our results are robust to structural shift consideration.

INSERT TABLE 6 HERE

Next, we control structural/policy measures as additional main effects to avoid a potential misspecification.[25] In fact, a number of empirical exchange rate regression models in the literature include some of these variables as additional controls.[26] Now the regression model becomes

$$REER_{it} = \alpha_1 + \beta_1 RCP_{it} + \sum_{s=-p}^{p} \delta_s \Delta RCP_{it+s} + RCP_{it}\boldsymbol{X}_{it}\boldsymbol{\beta}_2 + X_{it}\boldsymbol{\beta}_3 + \theta a_i + u_{it} \tag{27}$$

Table 7 shows the estimation results for the long-run elasticity and interaction effects with additional controls. After including structural/policy variables as independent control factors, we have to be extremely cautious in interpreting the results. For example, conditional on international reserves for a selected country group with high stock of international reserves, the real exchange rate response tends to be amplified to commodity price shocks. However, we find that the overall influence of *RES* on *REER* is negative.[27] Our finding confirms that international reserves play a mitigating role for *REER* in the face of commodity terms-of-trade shocks, consistent with the empirical results in Aizenman et al. (2012).

The last column (5) is reported to check the robustness of our specification by including all the interaction terms and control variables together. Overall, the result confirms our prior findings. Nonetheless, coefficients of globalization measures become weak in terms of the magnitude and statistical significance. The potential presence of high multicollinearity is one reason for this result. Because, being dependent heavily on commodity exports, a major commodity producer's value of trade is likely to be highly correlated with the world price of its commodity exports.

[25] Our model specification may be subject to the criticism by not including the following standard determinants of long-run exchange rates such as sectoral productivity differentials, government spending, cumulated current account imbalances, and interest rate differentials amongst others. However, omission of those factors is not likely to cause sizeable bias because the *RCP* well satisfies the exogeneity assumption (Chen and Rogoff, 2003; Cashin et al., 2004). Consequently, by assumption, the omitted potential long-run exchange rate determinants are not correlated with our key explanatory variable *RCP*. In other words, the coefficient estimate of *RCP* is still unbiased.

[26] It is a common practice in the literature to include exchange rate regimes, international reserves (Oomes and Kalcheva, 2007) and openness measures (*FO*, *TO*: Aizenman et al., 2012, *TO*: MacDonald and Ricci, 2004) as a determinant of the real exchange rate.

[27] We reach this conclusion because β_3 (from equation (27)) outweighs β_2 for the vast majority of the observations of *RCP*.

INSERT TABLE 7 HERE

We also test the robustness of our main results by considering potential presence of cointegration between the real effective exchange rate and interaction terms. This may be a necessary step as interaction variables introduced in our long-run model are constructed by multiplying non-stationary commodity prices, which are cointegrated with the real effective exchange rate, and binary dummy structural/policy variables. Standard panel unit root and cointegration tests, which are the ones employed in section 3, detect the long-run cointegrating relationship between the real effective exchange rate and interaction terms involving openness and policy measures. As shown in Table 8, accounting for these additional cointegrating relationships by including contemporaneous, 1 lead and 1 lag of changes of cointegrated commodity price and interaction terms in each specification does not alter our main findings.

INSERT TABLE 8 HERE

Furthermore, we explore robustness of the long-run DOLS estimation results to longer leads and lags of commodity price term. This experiment, although not presented to save a space, generates little difference and validates our choice of lead and lag in dynamic OLS specification.

5 Concluding remarks

Unlike most of the literature that relies on the set of developed country time series data, we generalize the long-run commodity currency phenomenon in a large panel of commodity-dependent developing countries. In addition, we show that, for our data set, the commodity price-exchange rate connection is much weaker in the short-run than has generally been observed in prior literature based on a small set of developed countries.

Although the price of commodity exports has a strong influence on their real exchange rates in a group of commodity-exporting countries, the magnitude of the real exchange rate response varies substantially across countries. In our structural model, we find the factors that influence the reaction of the real exchange rate to commodity price shocks. These include the nation's export market structure, its monetary policy choices and its degree of openness. Our empirical results, which support the theoretical role of these factors in shaping the commodity price-exchange rate link, are summarized below.

The response of the real exchange rate to a commodity boom would be amplified if a country is characterized by any of the following traits: i) heavy commodity export dependency, ii) monopoly pricing power by possessing a dominant export share in world markets, iii) fixed nominal exchange rate, iv) inflation-targeting, v) low international reserves holdings, vi) open financial market and vii) low degree of trade dependence.

Facing a rising trend of commodity prices mainly driven by the strong global demand during the past decade, commodity-dependent countries are recently exposed to real appreciation pressure, which is likely to induce higher volatility in domestic output and price levels. Understanding the exchange rate response to fluctuations in commodity prices should be of particular interest to policy makers in commodity exporters as the conduct of monetary policy and globalization strategy can alter the magnitude of exchange rate responses and consequently incur the macroeconomic adjustment costs.

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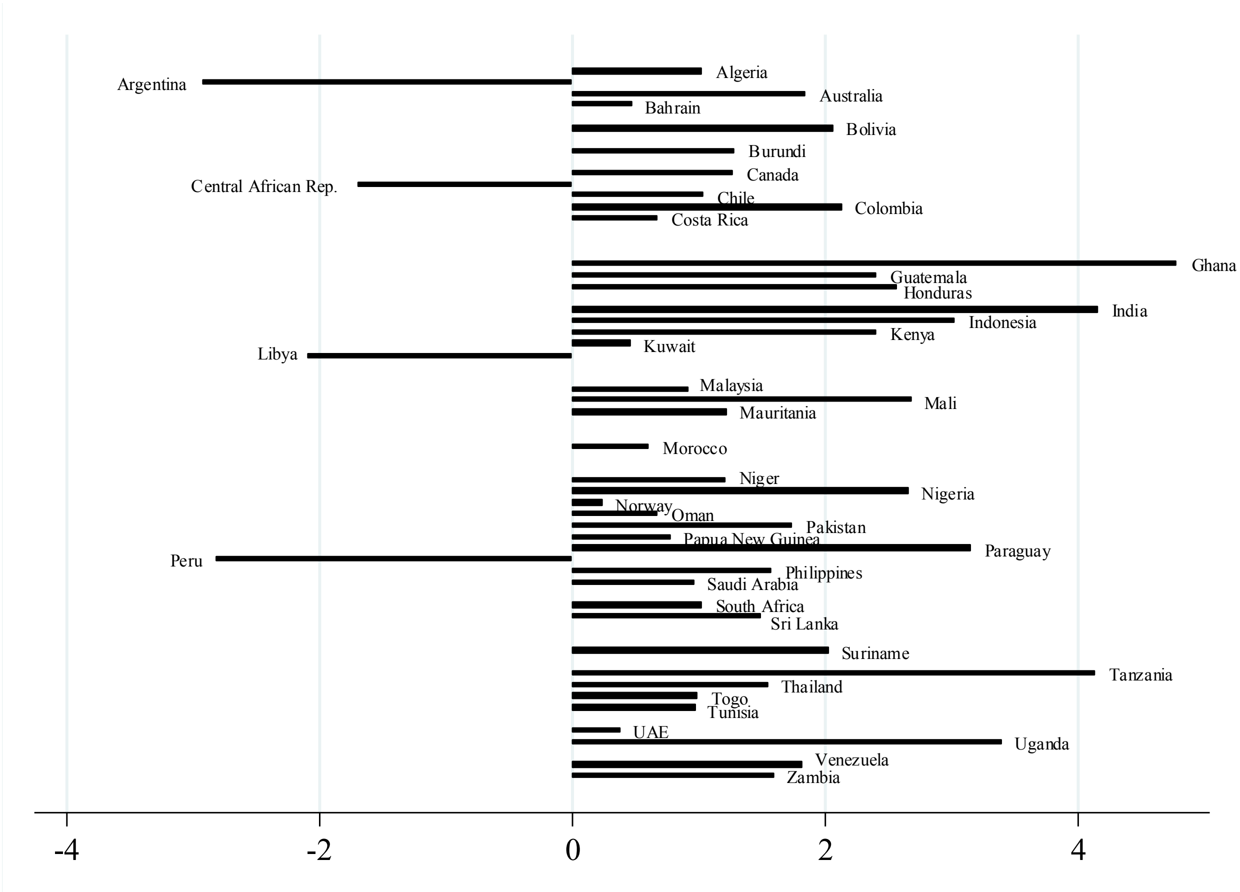
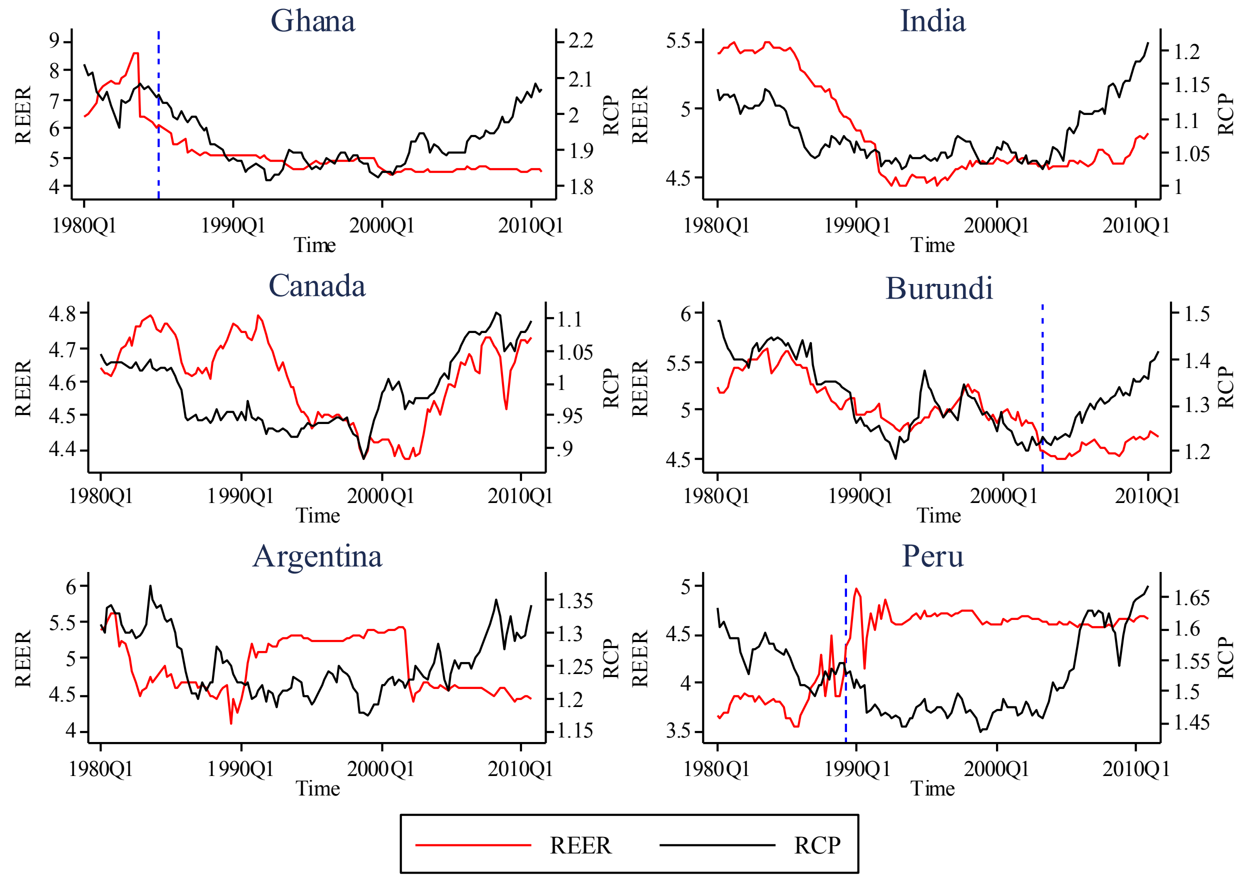
Figure 1. Distribution of commodity price elasticity of the real exchange rate across countries[28]

$$REER_t = \alpha_1 + \alpha_2 RCP_t + \varepsilon_t$$



[28] In this illustrative figure, we ignore the time series properties of real effective exchange rate ($REER$) and real commodity prices (RCP), and estimate the elasticity for each of the 66 countries in our sample using a simple regression model in levels: $REER_t = \alpha_1 + dt + \alpha_2 RCP_t + \varepsilon_t$, where $REER$ and RCP are in logarithm and t is a time trend. Newey-West standard errors are used, and we plot only the elasticity estimates that are statistically significant at the 5 percent level.

Figure 2. Time series plots of real effective exchange rate and real commodity prices[29]



[29] Vertical dashed lines in plots for Burundi, Ghana and Peru indicate the structural shift dates detected by the Gregory and Hansen (1996) test.

Table 1. Pesaran (2004)'s test of cross-sectional independence

CD test statistic	Average absolute value of the off-diagonal elements
121.22***	0.445

Note: CD test statistic is based on the residuals of the regression model $REER_{it} = \beta_0 + \beta_1 RCP_{it} + e_{it}$, where $REER$ and RCP are in logarithm. We include country-fixed effects. *** indicates rejection of the null hypothesis at the 1% significance level.

Table 2: Panel unit-root test results

Method	*REER*		*RCP*	
	Levels	1st differences	Levels	1st differences
LLC	3.520	-52.872***	8.615	-40.329***
	(0.999)	(<.01)	(1.00)	(<.01)
IPS	2.395	-48.216***	9.207	-51.619***
	(0.992)	(<.01)	(1.00)	(<.01)
CIPS	0.992	-7.447***	2.634	-11.718***
	(0.839)	(<.01)	(0.996)	(<.01)

Note: Reported are the t^*-statistic for LLC, W-stat for IPS, and Z[t-bar] statistic for CIPS test. The associated p-values of the test statistics are given in parentheses. *** indicates rejection of the null hypothesis at the 1% significance level. For the series in levels, we include individual trends and individual intercepts, while only country-specific intercepts are included for the series in first differences. Lag lengths are selected based on the modified Akaike Information Criterion. *REER* and *RCP* are in logarithm.

Table 3: Panel cointegration test results

Method			
Kao test			
ADF *t*-statistic	-1.894** (0.029)		
Pedroni test			
Within-dimension		*Between-dimension*	
Panel v-statistic	3.408*** (<.01)	Group ρ-statistic	-4.898*** (<.01)
Panel ρ-statistic	-7.357*** (<.01)	Group PP-statistic	-6.138*** (<.01)
Panel PP-statistic	-7.521*** (<.01)	Group ADF-statistic	-2.133** (0.017)
Panel ADF-statistic	-3.186*** (<.01)		
Westerlund test			
G_t	-2.274*** (<.01)		
G_a	-9.179*** (<.01)		
P_t	-14.762*** (<.01)		
P_a	-6.789*** (<.01)		

Note: Reported are the test statistics for each test. The corresponding p-values are given in parentheses. Statistical significance at the 1 and 5 percent levels is denoted by *** and **, respectively. For the Kao test, an individual intercept is included only, while the individual intercept and individual trend are included for the Pedroni (Engle-Granger based) test. Lag lengths are automatically selected based on the modified Akaike Information Criterion for both Kao and Pedroni tests. For Westerlund test, we use the Akaike Information Criterion to choose optimal lag and lead lengths for each series. Width of Bartlett-kernel window is set according to $4(T/100)^{(2/9)} \approx 4$. We allow for a constant but no deterministic trend in the cointegrating relationship.

Table 4. Estimation results for long-run elasticity and interaction effects: DOLS(1,1)

Dependent variable: $REER_t$					
	(1)	(2)	(3)	(4)	(5)
RCP_t	0.929**	1.023**	1.023***	0.588**	0.559***
	(0.415)	(0.410)	(0.367)	(0.228)	(0.198)
RCP_t*TO_t		-0.206***			-0.123***
		(0.028)			(0.016)
RCP_t*FO_t		0.043**			0.038**
		(0.017)			(0.015)
RCP_t*IT_t			-0.023		0.024
			(0.025)		(0.021)
RCP_t*EXR_t			0.047***		0.077***
			(0.016)		(0.015)
RCP_t*RES_t			-0.078***		-0.046**
			(0.02)		(0.019)
RCP_t*CEX_t				0.063***	0.058***
				(0.02)	(0.014)
RCP_t*MSH_t				0.032*	-0.0004
				(0.017)	(0.018)
Observations	7349	7108	6869	5946	5550
R^2	0.35	0.41	0.38	0.43	0.49
F-statistic		31.26***	30.59***	6.47***	14.53***

Note: DOLS(1,1) procedure includes contemporaneous, 1 lead and 1 lag of changes of cointegrated commodity price variable although they are suppressed to save a space. The specification also includes country fixed effects. Driscoll-Kraay standard errors are reported in parentheses. ***, **, * indicate significance at the 1%, 5%, and 10%, respectively.

Table 5. Estimation results for short-run elasticity and interaction effects: First difference

Dependent variable: $\Delta REER_t$

	(1)	(2)	(3)	(4)	(5)
ΔRCP_t	0.083*	0.093	0.208	0.205**	0.434**
	(0.048)	(0.115)	(0.131)	(0.079)	(0.209)
$\Delta RCP_t{*}TO_t$		-0.097			0.027
		(0.103)			(0.109)
$\Delta RCP_t{*}FO_t$		0.105			0.083
		(0.076)			(0.08)
$\Delta RCP_t{*}IT_t$			0.561**		0.43*
			(0.233)		(0.242)
$\Delta RCP_t{*}EXR_t$			-0.107		-0.232
			(0.115)		(0.142)
$\Delta RCP_t{*}RES_t$			-0.132		-0.133
			(0.082)		(0.084)
$\Delta RCP_t{*}CEX_t$				-0.17**	-0.192**
				(0.077)	(0.092)
$\Delta RCP_t{*}MSH_t$				-0.049	-0.15
				(0.085)	(0.1)
Observations	7463	7211	6973	6026	5624
R^2	0.004	0.005	0.01	0.01	0.01
F-statistic		1.47	2.55*	2.43*	2.58**

Note: Country fixed effects are included in all specifications. Driscoll-Kraay standard errors are reported in parentheses. ** and * indicate significance at the 5% and 10%, respectively.

Table 6. Robustness: Accounting for structural shifts

Dependent variable: $REER_t$

	(1)	(2)	(3)	(4)	(5)
RCP_t	0.930**	1.024**	1.021***	0.587**	0.559***
	(0.415)	(0.411)	(0.368)	(0.228)	(0.198)
RCP_t*TO_t		-0.206***			-0.123***
		(0.029)			(0.016)
RCP_t*FO_t		0.043**			0.037**
		(0.017)			(0.015)
RCP_t*IT_t			-0.023		0.024
			(0.025)		(0.021)
RCP_t*EXR_t			0.047***		0.077***
			(0.016)		(0.015)
RCP_t*RES_t			-0.079***		-0.046**
			(0.02)		(0.019)
RCP_t*CEX_t				0.063**	0.058***
				(0.020)	(0.014)
RCP_t*MSH_t				0.031*	-0.001
				(0.017)	(0.018)
Observations	7349	7108	6869	5946	5550
R^2	0.35	0.41	0.39	0.43	0.49
F-statistic		30.62***	30.61***	6.45***	14.45***

Note: See note of Table 4.

Table 7. Robustness: Controlling for conditional factors

Dependent variable: $REER_t$					
	(1)	(2)	(3)	(4)	(5)
RCP_t	0.608***	0.849**	0.879**	0.736***	0.549***
	(0.189)	(0.4)	(0.359)	(0.224)	(0.182)
RCP_t*TO_t		0.005			-0.051
		(0.066)			(0.049)
RCP_t*FO_t		0.164**			-0.011
		(0.063)			(0.063)
RCP_t*IT_t			0.178**		0.235***
			(0.071)		(0.066)
RCP_t*EXR_t			0.041		0.149***
			(0.035)		(0.047)
RCP_t*RES_t			0.2***		0.219***
			(0.037)		(0.034)
RCP_t*CEX_t				-0.116*	-0.124*
				(0.069)	(0.063)
RCP_t*MSH_t				-0.074	-0.055
				(0.079)	(0.076)
TO_t	-0.146***	-0.317***			-0.076
	(0.021)	(0.085)			(0.068)
FO_t	0.044**	-0.160**			0.063
	(0.017)	(0.076)			(0.071)
IT_t	0.018		-0.235***		-0.268***
	(0.023)		(0.066)		(0.068)
EXR_t	0.064***		-0.012		-0.12**
	(0.016)		(0.055)		(0.059)
RES_t	-0.087***		-0.397***		-0.355***
	(0.025)		(0.061)		(0.051)
CEX_t	0.067***			0.221**	0.209***
	(0.017)			(0.097)	(0.076)
MSH_t	0.01			0.127	0.06
	(0.021)			(0.089)	(0.083)
Observations	5550	7108	6869	5946	5550
R^2	0.49	0.42	0.41	0.44	0.52
F-statistic		3.96**	10.68***	2.07	11.41***

Note: See note of Table 4.

Table 8. Robustness: Accounting for additional cointegrating relationships

Dependent variable: $REER_t$				
	(1)	(2)	(3)	(4)
RCP_t	0.929**	1.046**	1.028***	1.041***
	(0.415)	(0.404)	(0.36)	(0.343)
RCP_t*TO_t		-0.244***		-0.2***
		(0.031)		(0.024)
RCP_t*FO_t		0.045**		0.039**
		(0.019)		(0.018)
RCP_t*IT_t			-0.023	0.017
			(0.026)	(0.023)
RCP_t*EXR_t			0.053***	0.059***
			(0.016)	(0.017)
RCP_t*RES_t			-0.086***	-0.083***
			(0.022)	(0.022)
Observations	7349	7081	6839	6716
R^2	0.35	0.42	0.39	0.44
F-statistic		35.69***	30.19***	27.56***

Note: DOLS(1,1) procedure includes contemporaneous, 1 lead and 1 lag of changes of cointegrated commodity price and interaction terms although they are suppressed to save a space. The specification also includes country fixed effects. Driscoll-Kraay standard errors are reported in parentheses. ***, **, * indicate significance at the 1%, 5%, and 10%, respectively.

Appendix

1. Choice of countries

We start from the set of countries chosen by Cashin et al. (2004) and modify the sample following Coudert et al. (2008). Ecuador and Nicaragua were excluded from our sample because of its dollarization that began in 2001 and unusual 1000% appreciation at the beginning of the sample period, respectively. Myanmar and Zimbabwe were dropped as well due to the unavailability of continuous time series data for commodity exports in UN COMTRADE and hyperinflation during the significant part of sample period (since 2002) that could distort an appropriate measure of exchange rate, respectively. The following nine major oil exporters were added: Algeria, Bahrain, Kuwait, Libya, Nigeria, Oman, Saudi Arabia, United Arab Emirates, and Venezuela. This procedure leaves a total of 63 commodity exporters including both non-oil and oil exporters. The majority of countries in our sample are developing economies (58 countries). The full list of countries is available in Table A1 in Appendix.

2. Real effective exchange rate (*REER*) and real commodity price index (*RCP*)

Following the literature in this line of research, we obtain real effective exchange rate (*REER*), an average of the bilateral real exchange rates between the country and its trading partners weighted by the respective trade shares of each trading partner, from the International Monetary Fund's (IMF) International Financial Statistics (IFS) and Information Notice System (INS). From its definition, an increase in real effective exchange rates implies a real appreciation of the domestic currency.

We define a real commodity price index as the world (nominal) price of country's commodity exports relative to the world price of manufactured goods exports. It is a common practice to measure the commodity terms-of-trade of countries with high commodity export dependence in this way in the literature (see Grilli and Yang, 1988; Deaton and Miller, 1996; Cashin et al., 2000). The annual commodity trade data are taken from the UN COMTRADE and the monthly world commodity price data are from the IMF Primary Commodity Prices and the World Bank Pink Sheet. We construct monthly commodity price indices using 58 commodities for 63 commodity exporters.[30] So, for each country i, commodity j, and time t, a country specific index of nominal commodity price is defined as

$$\ln NCP_{it} = \sum_{j=1}^{J} W_{ij} (\ln P_{jt}),$$

$$W_{ij} = \frac{1}{T}\sum_{t=1}^{T} X_{ijt} \Big/ \frac{1}{T}\sum_{t=1}^{T} CX_{it}$$

where X is an export volume of individual commodity j and CX is a total commodity export volume. The weights (W) remain constant over time in order to eliminate the quantity effect from the price measure. This definition is similar to Cashin et al. (2004) but we use the period-average export values of each commodity and total commodities between 1980 and 2010. Commodity prices are expressed in real terms (*RCP*) through deflation by the IMF's unit value index of manufactured exports (MUV) of industrial economies. Note that throughout the paper, both *REER* and *RCP* are in log forms.

[30] We try to include as many commodities as possible if their prices are available in the IMF Primary Commodity Prices and World Bank Pink Sheet. Platinum, plywood and steel are excluded because we have no information about the corresponding SITC codes.