What Makes a Commodity Currency?*

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This draft: September, 2014

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 panel data analysis using 63 countries for 1980-2010 finds that, in accordance with theory, the long-run cointegrating relationship between the real exchange rate and commodity export prices depends on the nation's export market structure, monetary policy choices and degree of trade and financial openness. We also show that the commodity price-exchange rate connection is much weaker in the short-run and for a group of oil-exporting countries. Given concerns for the Dutch disease or resource curse, our findings are of particular relevance for monetary policy-making and for globalization strategy in commodity-exporting developing economies.

JEL classification: C32; C33; F31; F41; O13
Keywords: Real exchange rate; Commodity prices; Panel cointegration; Commodity exports; Exchange rate regime; International reserves

* First draft: November, 2012. We are grateful to Karen Lewis, James D. Hamilton, Stephen Turnovsky, Seik Kim, Sebastian Fossati, Kar-yiu Wong, Eric Zivot, Chang-Jin Kim, and seminar participants at the University of Washington, University of California, Riverside, Bank of Korea, Academia Sinica, the 2013 NBER Economics of Commodity Markets Meeting, the 2014 Society for Nonlinear Dynamics and Econometrics Symposium, and the 2014 Society for Computational Economics-Computing in Economics and Finance Conference for their helpful comments and suggestions on earlier drafts. We also thank Paul Cashin and Susan Yang at the IMF for providing some of the data used in this paper. All remaining errors are our own.
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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 their 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 real currency appreciation can crowd out the exports of non-commodity industries by undermining their price competitiveness in the world trade. 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 from diverse perspectives. First, the paper explores the intermediate role of structural and policy factors in determining the strength of real exchange rate-commodity price connection. Second, the paper makes a clear distinction of workings of commodity currencies between in the short- and long-run, which is often neglected in the literature. Lastly, the paper documents differences in commodity currencies and oil currencies in terms of their responses to a commodity/oil price shock.

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[1] Currencies that respond significantly to the world prices of their corresponding country's commodity exports are called "*commodity currencies*". See Edwards (1986), 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. Ricci et al. (2008), Coudert et al. (2008), and Bodart et al. (2011, 2012) study commodity currencies based on a group of commodity-dependent economies using a panel data approach.
[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) finds 19 commodity currencies with the estimates of long-run commodity price elasticity of real exchange rate ranging from 0.16 for Iceland to 2.03 for Ecuador.

As a preview, Figure 1 illustrates the large heterogeneity in the domestic currency responses to movements of world commodity prices across major commodity-exporting countries, whose export earnings in primary commodity products generally account for more than half of total export earnings.[4] Regressing country-by-country the real effective exchange rate (*REER*) on the country-specific real commodity price index (*RCP*), we find 43 countries out of 63 to have a statistically significant commodity price coefficient at the 5 percent level.[5] Elasticity estimates range from −4.96 (Libya) to 7.63 (Ghana) with a median value of 1.[6]

INSERT FIGURE 1 HERE

What may account for this heterogeneity? To answer this question, we first present a small-open economy, traded/non-traded goods model in the next section. Our model suggests that three groups of factors affect the link between *REER* and *RCP*: the nation's degree of openness (both trade and financial), monetary policy choices (in the form of inflation-targeting, nominal exchange rate flexibility and international reserves management) and export market characteristics (i.e., degree of commodity export dependency and export share in world markets). Our empirical results broadly support this theoretical view. More specifically, the long-run reaction of the real exchange rate to a commodity boom would be larger if a country is characterized by any of the following traits: i) open financial market, ii) low degree of trade openness, iii) fixed nominal exchange rate, iv) low level of international reserves, v) heavy commodity export dependency, and vi) dominant global commodity production share. Furthermore, our estimation results demonstrate a strong long-run *REER-RCP* connection,

[4] We focus on commodity exporters not importers because the theoretical channel we discuss in the next session relies on the link between the price of exportable commodities and non-traded goods. In commodity importing economies, imported commodities are typically used as intermediate inputs to produce final outputs. For this reason, without knowing commodity input content of various sectors, it is unclear ex ante how a commodity price shock is transmitted into the commodity-importing country's overall price level.

[5] 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. We interpret an increase in the real effective exchange rate 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. Note that *REER* and *RCP* are in logarithm in all of our empirical procedures. More information about the *REER* and *RCP* including their construction and data sources is presented in Appendix.

[6] Reported median here is from a distribution including both short- and long-run elasticity estimates. Both elasticity estimates for a full sample appear in Table A1 in Appendix. Note that the median would be 1.76 if we consider the long-run elasticity estimates only. The short-run elasticity estimates have a narrower distribution ranging from -0.63 (Venezuela) to 2.44 (Brazil) with a median value of 0.67.

generalizing the commodity currency phenomenon in a large group of developing countries. However, in contrast to previous studies based on the currencies of a small set of developed countries, we find much weaker evidence of the commodity currency phenomenon in the short-run.[7] We also find a weaker *REER-RCP* relation for a group of oil-exporting countries than the non-oil commodity counterparts.

During our sample period between 1980 and 2010, many developing countries in the world experienced a significant structural change in policies including exchange rate reforms and the adoption of inflation targeting. A series of currency crises also have affected macroeconomic conditions of this country group. All of these have a potential to affect the level of the real exchange rate and consequently a relationship between *REER* and *RCP*. We thus check the sensitivity of our key results and find they are robust to structural shift consideration.

While high commodity prices of any type bring about higher export revenue for a country exporting that commodity, they 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. Therefore, effectively managing these adverse consequences of commodity price fluctuations is a natural interest of policy makers in commodity exporting countries. Results in this paper help them to find appropriate policy responses to stabilize their economy by effectively dampening rather than amplifying the costly 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 estimation results and their robustness using a non-stationary panel data set. Section 5 concludes.

[7] Chen et al. (2010) exploits short-run asset pricing dynamics, focusing on five developed economies including Australia, Canada, New Zealand, South Africa and Chile.

2 Determinants of commodity price elasticity

In this section, we present a theoretical framework that highlights the impact of commodity price shocks on the real exchange rate in a commodity exporting country. The model allows us to discuss the effect of economic determinants such as a country's structural factors and policy choices on the strength of exchange rate-commodity price connection.

2.1 The baseline model

2.1.1 Production

Consider a small open economy that produces three types of goods, exportable commodities (*XC*), exportable manufactures (*XM*) and non-traded (*N*) in a competitive market. With a constant-returns-to-scale technology, production functions in each sector are given by

$$Y_{XC} = A_{XC} L_{XC}^{\alpha_{xc}} H_{XC}^{\beta_{xc}} K_{XC}^{1-\alpha_{xc}-\beta_{xc}} \tag{1}$$

$$Y_{XM} = A_{XM} L_{XM}^{\alpha_{xm}} H_{XM}^{\beta_{xm}} K_{XM}^{1-\alpha_{xm}-\beta_{xm}} \tag{2}$$

$$Y_N = A_N L_N^{\alpha_n} K_N^{1-\alpha_n} \tag{3}$$

where A_i, L_i, H_i and K_i are productivity shock, unskilled labor, skilled labor, and capital stock used in sectors i = *XC*, *XM*, *N* respectively; α_i and β_i are a share of unskilled and skilled labor in production process in sector i. Note that the skilled labor (*H*) is necessary to produce exportable goods, a part of which is consumed by domestic residents. Capital is allowed to move between sectors and countries, but labor (both skilled and unskilled) is assumed to migrate between sectors only within the country. The total domestic labor supply is inelastically given by $L = L_{XC} + L_{XM} + L_N$ and $H = H_{XC} + H_{XM}$. 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_L and w_H) equalization across sectors.

In our model, we take exportable manufactures (*XM*) as the numeraire, with P_{XC} the world price of exportable commodities exogeneously given to the small open economy and P_N the domestic price of non-traded goods, both measured in terms of exportable manufactures. The law of one price holds for the exportable goods so that

$$EP_i = P_i^* \qquad \text{for } i = XC, XM \tag{4}$$

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. Profit-maximizing optimal allocation of factors in both exportable goods is given by

$$P_i \alpha_i A_i h_i^{\beta_i} k_i^{1-\alpha_i-\beta_i} = w_L \tag{5}$$

$$P_i \beta_i A_i h_i^{\beta_i - 1} k_i^{1-\alpha_i-\beta_i} = w_H \tag{6}$$

$$P_i (1 - \alpha_i - \beta_i) A_i h_i^{\beta_i} k_i^{-\alpha_i-\beta_i} = r \tag{7}$$

where $h_i \equiv H_i/L_i$ and $k_i \equiv K_i/L_i$ stand for the skilled-unskilled labor ratio and capital-unskilled labor ratio respectively in sectors $i = XC$, XM, and $P_{XM} = 1$. By combining the above first-order conditions, we derive the following zero-profit conditions for each exportable sector:

$$P_{XC} A_{XC} h_{XC}^{\beta_{xc}} k_{XC}^{1-\alpha_{xc}-\beta_{xc}} = w_L + w_H h_{XC} + r k_{XC} \tag{8}$$

$$A_{XM} h_{XM}^{\beta_{xm}} k_{XM}^{1-\alpha_{xm}-\beta_{xm}} = w_L + w_H h_{XM} + r k_{XM} \tag{9}$$

To simplify further, we assume both exportable sectors share a common rate of productivity shock $\hat{A}_{XC} = \hat{A}_{XM} = \hat{A}_X$ and have the same share of capital income in output $\mu_{K,XC} = \mu_{K,XM}$ where $\mu_{K,i} \equiv rK_i/P_iY_i$ for sectors $i = XC$, XM.[8] In addition, let $\mu_{H,i} \equiv w_H H_i/P_i Y_i$ and $\mu_{L,i} \equiv w_L L_i/P_i Y_i$ be skilled and unskilled labor's share of the income for a given sector i. Under these assumptions, taking a log-differentiation of equations (8) and (9) making use of (6) and (7) yields:

$$\widehat{w}_L = \frac{\varphi \hat{P}_{XC} + \hat{A}_{XC}}{\mu_{L,XC} + \mu_{H,XC}} \tag{10}$$

$$\text{where } \varphi = \frac{\mu_{H,XM}}{\mu_{H,XM} - \mu_{H,XC}}.$$

[8] A hat above the variable denotes a logarithmic derivative.

On the other hand, from the zero-profit condition in the non-traded good sector, we obtain

$$\hat{P}_N = \mu_{L,N}\hat{w}_L - \hat{A}_N \tag{11}$$

Combining equations (10) and (11) gives

$$\hat{P}_N = \frac{\mu_{L,N}}{\mu_{L,XC} + \mu_{H,XC}}\left(\varphi\hat{P}_{XC} + \hat{A}_{XC}\right) - \hat{A}_N \tag{12}$$

Empirically, non-traded goods tend to be more labor-intensive than exportable goods, implying that $\mu_{L,N} \geq \mu_{L,XC} + \mu_{H,XC}$. Moreover, if the production of the exportable manufacturers is relatively more skilled-labor intensive than exportable commodities, $\varphi > 0$. It then follows that the price of non-traded (P_N) is a positive function of the price of exportable commodities (P_{XC}). Equation (12) also shows the standard Balassa-Samuelson effect: a higher productivity in the commodity sector leads to an increase in the price of the non-traded through the wage adjustment.[9]

2.1.2 Real exchange rate determination based on CPI

Our model economy is inhabited by a continuum of identical individuals that supply production inputs and consume four types of goods – non-traded (*N*), imported (*M*), exportable commodities (*XC*) and exportable manufactures (*XM*). Let's assume that home residents derive utility by consuming θ_N share of non-traded, θ_M share of imported, θ_{XC} share of exportable commodities and $(1 - \theta_N - \theta_M - \theta_{XC})$ share of exportable manufactures, where the law of one price holds for all types of tradable goods. Then a representative consumer's utility function takes the following Cobb-Douglas form:

$$U = C_N^{\theta_N} C_M^{\theta_M} C_{XC}^{\theta_{XC}} C_{XM}^{1-\theta_N-\theta_M-\theta_{XC}} \tag{13}$$

[9] Note that with perfect international capital mobility and free movements of labor (skilled and unskilled) across sectors, the relative price of non-traded is entirely determined by the production side of our model and is independent of demand side factors.

where C_i is consumption of good $i = N, M, XC, XM$. Using the consumption-based price index for home and foreign economies and the relation derived in equation (12), we can write the real exchange rate (Q), the relative price of domestic consumption basket in terms of foreign consumption basket, as follows:[10]

$$Q = \frac{EP}{P^*} = \frac{EP_N^{\theta_N} P_M^{\theta_M} P_{XC}^{\theta_{XC}}}{(P_N^*)^{\theta_N^*}(P_M^*)^{\theta_M^*}(P_{XM}^*)^{1-\theta_N^*-\theta_M^*}} = \left(E\frac{P_N}{P_N^*}\right)^{\theta_N} P_{XC}^{\theta_{XC}} = \left(E\frac{f(\overset{(+)}{P_{XC}})}{P_N^*}\right)^{\theta_N} P_{XC}^{\theta_{XC}} \qquad (14)$$

where we assume $\theta_N = \theta_N^*$ and $\theta_M = \theta_M^*$, and normalize the home price of exportable manufactures to one. The equation (14) shows that, given P_N^*, the real exchange rate in the home country appreciates in response to an increase in the price of commodity exports, with the extent of this appreciation depending on the variables/parameters present in the equation. Detail discussions about these relevant variables/parameters follow in the next section.

2.2 Factors influencing the commodity price elasticity

2.2.1 Degree of openness

Trade Openness (*TO*). The equation (14) shows that the elasticity of the real exchange rate with respect to the price of exportable commodities depends on θ_N, which captures a share of non-traded in a basket of domestic consumption. Hence, holding all other things constant, if the economy's consumption depends heavily on non-traded goods with a large θ_N, the real exchange rate response to an increase in the price of exportable commodities would be relatively large. In other words, a country with a high degree of trade openness by having a relatively small share of non-traded is likely to have a lower commodity price elasticity of the real exchange rate.

[10] Similar to Cashin et al. (2004), we assume that the foreign economy, which is a *trade partner* of the home country, produces intermediate (*I*), non-traded (*N*), and final exportable (*M*) goods using labor (unskilled and skilled) and capital, and consumes θ_N^*, θ_M^* and $(1-\theta_N^*-\theta_M^*)$ shares of the non-traded, final exportable and imported manufactures (*XM*), respectively. The foreign country's final exportable good is a manufactured good, a fraction of which is exported to the home country and consumed by the home residents. We assume that the foreign firms use *v* share of intermediate goods and (1 - *v*) share of commodities imported from the home country to produce the final exportable goods but a commodity input share is small enough that world price changes in commodities have the negligible effect on CPI of the foreign economy.

Financial Openness (*FO*). A commodity price boom is expected to attract foreign capital to the exportable commodity sector, raising marginal productivity of both unskilled and skilled labor and wages by equations (5) and (6). Through the channels described in equations (10) and (11), this should generate an increase in wage of unskilled workers and eventually the price of non-traded, resulting in an equilibrium real exchange rate appreciation. Apparently, capital flows play a role in explaining the Balassa-Samuelson channel. Let's now introduce financial market frictions in the form of restricted capital inflow to a commodity-exporting country. In this case, the real appreciation pressure caused by a rise in commodity price would be relatively small as improvement of the marginal productivity of labor in exportable commodities is likely to be less pronounced with inelastic supply of capital. This suggests that a stronger real exchange rate-commodity price connection may exist in a county with a higher degree of financial openness.

2.2.2 Monetary/exchange rate policy choices

Inflation Targeting (*IT*) and *Exchange Rate Regime* (*EXR*). Monetary/exchange rate policy options designed to control the movements of nominal exchange rate (E) or domestic price level (P) are also important candidates influencing the link between the real exchange rate and commodity prices as shown in equation (14). Under the perfect cross-border capital mobility, however, a monetary authority faces a trade-off between fixing the exchange rate and keeping an autonomous monetary policy instrument.[11] Recognizing this interdependent nature of monetary/exchange rate policy choices, we interpret their role in affecting a commodity price elasticity together as a group rather than individually. On the other hand, although our focus in this section is mainly in the long-run channel, we may get different results for the effect of monetary policy choices between in the long- and short-run due to the existence of nominal rigidity. In theory, it is unclear whether adopting an inflation targeting regime or a currency peg is more effective in stabilizing the real exchange rate in response to a commodity price boom.[12]

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

[12] While we admit that a monetary policy should be neutral in the long-run, it may have a long-run effect in practice as massive international reserve holdings allow a country to pursue both a high level of exchange rate stability and a relatively high weighted average of the other two trilemma policy objectives. Also note that, although our model emphasizes a transmission of commodity price shocks to the real exchange rate only, 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.

This is because we do not know a priori how much the domestic price level in a commodity-exporting country would adjust relative to the nominal exchange rate when the commodity price rises. Empirical evidence in the literature is also mixed. For example, 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. On the contrary, Bodart et al. (2011) finds that a flexible exchange rate regime tends to decrease the effect of commodity price shocks on the real exchange rate.

International Reserves (*RES*). A sizeable stock of international reserves can help stabilize the exchange rate by providing a country's foreign exchange market with extra liquidity. Commodity-exporting countries are not an exception and can also benefit from international reserves. In fact, Aizenman et al. (2012) shows that large reserves effectively lower the volatility of the real exchange rates in commodity-dependent Latin American countries that frequently face commodity terms-of-trade shocks. Furthermore, in a country with an open capital market, the effectiveness of monetary and exchange rate policies can depend on the size of international reserves because hoarding large reserves may relax the policy constraints under an open-economy trilemma. Aizenman et al. (2013) supports this view by documenting the presence of loose compatibility amongst three objectives, namely exchange rate stabilization, capital mobility, and domestic monetary autonomy, in emerging markets with ample international reserves.

2.2.3 Export market characteristics

Commodity Export Dependency (*CEX*). The equation (14) shows that, holding everything else constant, the real exchange rate response to a commodity price change is larger with a higher share of exportable commodities θ_{XC} in a domestic consumption basket. Therefore, the real exchange rate of a country with high commodity export dependency is expected to be more sensitive to commodity price shocks than otherwise.

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 share of the global commodity production and, as a result, has some degree of market power.[13] Consider a country that has a monopoly power in the world market for a commodity in the sense that a large volume of its exports places downward pressure on the world price of the commodity. Domestic producers in such a country would expand the production and export more to increase their revenue if a domestic currency depreciates. As a result, the world price of the commodity would fall due to the large supply. Based on this logic, we can write the production of exportable commodities as a negative function of the real exchange rate:

$$Y_{XC} = Y_{XC}\left(\overset{(-)}{Q}\right) \tag{15}$$

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:

$$P_{XC} = P_{XC}\left(\overset{(-)}{Y_{xc}}\right) \tag{16}$$

holding everything else that possibly influences the commodity price constant. Thus, from equations (14) 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 this reduction of supply pushes up the world commodity price further by (16) as, by assumption, the country is large. This logic suggests that the exchange rate-commodity price connection can be stronger when the commodity prices are endogenously determined by a country with the sufficiently large market power.

Finally, the theoretical effect of factors influencing the commodity price elasticity of real exchange rate $\left(\frac{\hat{Q}}{\hat{P}_{XC}}\right)$ discussed in this section can be summarized as follows:

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[13] Examples of countries with a sufficiently large market share of commodity exports include Chile (copper), Cote d'Ivoire (cocoa), Malaysia (palm oil), and Philippines (coconut oil). Each of these countries often accounts for more than one third of world production of its primary commodity.

$$\frac{\hat{Q}}{\hat{P}_{XC}} = g\left(\overset{-}{TO}, \overset{+}{FO}, \overset{+/-}{IT}, \overset{+/-}{EXR}, \overset{-}{RES}, \overset{+}{CEX}, \overset{+}{MSH}\right) \tag{17}$$

where the signs above the variables indicate the expected effect of these variables on the commodity price elasticity.

3 Empirical procedures

3.1 Baseline regression model

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

$$REER_t = \alpha_0 + \alpha_1 RCP_t + \varepsilon_t \tag{18}$$

where $t = 1,\ldots,T$ indexes the time-series, $REER_t$ 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_1. Our goal is to identify factors that may account for a large variation of α_1's across countries. In other words, we want to explain the parameter α_1 using a set of country-specific variables X such that

$$\hat{\alpha}_{1i} = h(X_i) = \beta_1 + X_i\boldsymbol{\beta_2} \tag{19}$$

where i indexes cross-sectional units, X includes seven factors of commodity price elasticity discussed in the previous section, and $\boldsymbol{\beta}_2$ is a vector of coefficients. Combining (18) and (19) under the exogeneity assumption, our empirical model is given by the following:[14]

$$REER_{it} = \alpha_0 + \beta_1 RCP_{it} + RCP_{it}X_{it}\boldsymbol{\beta_2} + u_{it} \tag{20}$$

[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 mathematically unjustifiable. Nevertheless, we present these steps here to show our motivation for the empirical estimation strategy.

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

3.2 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. The choice of sample countries and definitions and sources of variables are presented in Appendix.[15] A majority of our control variables are available only at an annual frequency and interpolated to a quarterly frequency through the "constant-match average". For our key variables *REER* and *RCP*, which are available at a monthly frequency, we use the "last observation" method that sets the quarterly observation equal to the value in the last of the corresponding monthly observations.

As a preliminary step, we show in Figure 2 the time-series of the *REER* and *RCP* using a small set of countries from our sample. Two developed (Australia and Canada) and two developing (Ghana and Peru) commodity-exporting countries are selected.[16] 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 possibility of having 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 in selected countries, except for Peru. Moreover, the relationship between the *REER* and *RCP* exhibits structural shifts in countries such as Ghana and Peru.[17] Selected shift dates are largely consistent with an economic event of a country in that period. For example, the *REER* of Ghana experienced a steep depreciation from the period after the Structural Adjustment Programme (SAP) in 1983, which

[15] Table A2 in Appendix contains descriptive statistics for the sample data used in our empirical estimation.
[16] See Table A3 in Appendix to learn major commodities exported by the selected four countries (and the rest of countries in our sample) and their share in aggregate commodity exports.
[17] Structural shift dates are indicated in Figure 2 by dashed vertical lines and reported in column 9 of Table A1 in Appendix. Gregory-Hansen (1996) cointegration test is used to locate regime shift dates and is discussed in detail in section 4.3.

included exchange rate reforms until 1990. Peru, on the other hand, experienced a dramatic appreciation of its domestic currency because of the hyperinflation episodes in the late 1980's.

INSERT FIGURE 2 HERE

In addition to these time series properties, cross-sectional dependence is likely to be important and present in our case 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. Driscoll and Kraay (1998) note that standard error estimates of commonly applied co-variance matrix estimation techniques such as OLS, White and Rogers are biased by erroneously ignoring spatial correlation in panel regressions and hence statistical inference that is based on such standard errors is invalid. Typically, ignoring cross-sectional dependence leads to overly optimistic standard error estimates. We thus conduct Pesaran (2004)'s cross-sectional dependence (CD) test and Table 1-a shows the test results. The null of cross-sectional independence is rejected at the 1 percent significance level, indicating that the regression residuals are cross-sectional dependent.

INSERT TABLE 1 HERE

Next, we check the null of unit-root for the *REER* and *RCP* by running the Levin-Lin-Chu (2002) (LLC) test, Im-Pesaran-Shin (2003) (IPS) test, and Pesaran's (2007) cross-sectionally augmented IPS (CIPS) test. Kao (1999), Pedroni (2004), and Westerlund (2007) panel conintegration tests allow us to test the null of no cointegration.[18] As evident from Table 1-b, we cannot reject the null of unit-root for *REER* and *RCP* at their levels and both series are integrated of order one. Table 1-c shows the results in favor of cointegration, suggesting the existence of long-run relation between *REER* and *RCP*.

[18] A theoretical background for each of panel non-stationarity and cointegration tests can be found in Baltagi (1995). See Chapter 12 of his book and references therein for more details.

3.3 Additional empirical specifications

We estimate equation (20) using the (country) fixed effect model to reduce the omitted variable bias caused by unobserved country-specific factors. Also, 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. Lastly, all structural/policy factors are converted into binary dummy variables using the sample median as a threshold value for each series.[19]

3.4 Long-run vs. short-run estimation strategies

3.4.1 Long-run estimation: Dynamic OLS

Recognizing non-stationarity and the presence of cointegration for *REER* and *RCP*, we apply DOLS (Dynamic Ordinary Least Squares), extended to a panel data analysis by Kao and Chiang (2000) and Mark and Sul (2003), to estimate the cointegrating parameter. The DOLS procedure brings contemporaneous, leads and lags of changes of cointegrated regressor to remove their deleterious short-run dynamic effects on the estimation of long-run cointegrating vector.[20] For a country i and time period t, the long-run estimation is carried out based on the following regression model:

$$REER_{it} = \alpha_i + \beta_1 RCP_{it} + \sum_{j=-p}^{p} \Delta RCP_{i,t+j}\gamma_j + RCP_{it}X_{it}\boldsymbol{\beta_2} + u_{it} \tag{21}$$

[19] For example, we set *CEX* = 1 if *CEX* > median(*CEX*) and *CEX* = 0 otherwise. The same rule applies to *FO*, *RES*, and *MSH* variables. For a trade openness measure (*TO*), we use the threshold set in Aizenman et al. (2012): A country is highly trade-dependent when a ratio of trade (= *EX* + *IM*) to 2×*NGDP* is greater than 0.3. For an exchange rate regime measure, our binary *EXR* takes a unity if IRR (2010)'s coarse classification code is equal to 1 or 2 (peg).

[20] An alternative methodology widely used in the panel analysis with non-stationary data is FMOLS (Fully Modified Ordinary Least Squares). Kao and Chiang (2000) compares the performance of panel FMOLS and DOLS and reports that the DOLS is superior in removing a finite sample bias. Note also that FMOLS requires a balanced panel and thus the estimation has to rely on a substantially reduced sample size. For these reasons, we adopt the DOLS procedure as our main long-run estimation method.

where α_i captures the country-fixed effect, β_1 is the long-run cointegrating coefficient, γ_j is a coefficient vector of leads and lags of the changes in real commodity price index, X_{it} is a set of structural/policy variables, $\boldsymbol{\beta_2}$ is a vector of coefficients of interaction terms and u_{it} is the disturbance term.

3.4.2 Short-run estimation: Error Correction Model

Under the cointegration setting, we use a simple error correction model (ECM) that allows separating short-term from long-term effects. The model takes the following form:

$$\Delta REER_{it} = \alpha_i + \sum_{j=1}^{p} \Delta REER_{i,t-j}\varphi_j + \sum_{j=0}^{p} \Delta RCP_{i,t-j}\delta_j + (\Delta RCP_{it})X_{it}\boldsymbol{\theta_2} + \lambda EC_{i,t-1} + \varepsilon_{it} \quad (22)$$

where the error correction term is defined as $EC_{i,t-1} = REER_{i,t-1} - \alpha_i - \hat{\beta}_1 RCP_{i,t-1}$ with $\hat{\beta}_1$ being the cointegrating parameter estimate, and φ_j , δ_j , $\boldsymbol{\theta_2}$ and λ are relevant regression coefficients.

4 What makes a commodity currency?

4.1 Commodity currency in the long-run

4.1.1 Main results

First of all, we note that there is a strong and robust link between *REER* and *RCP* in the long-run across the different specifications from columns (1) to (5) in Table 2. Indeed, there is almost one-for-one cointegrating relation between them and this relation is statistically significant. For example, one percent permanent increase in commodity price index will cause the real effective exchange rate to appreciate by 0.93 percent according to the specification (1). The effect of globalization is presented in column (2). A higher trade dependency 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 theoretical prediction of section 2. For policy variables in column (3), we find the response of *REER* to the *RCP* fluctuations tends to be larger under a peg and smaller with large international reserve holdings. The results in column (3) suggest that a flexible exchange rate regime better insulate the economy from commodity price shocks by stabilizing the real exchange rate in the long-run. Our results also confirm the buffering role of foreign reserves by mitigating the impact of external shocks on the real exchange rate. Results in column (4) show that the greater the commodity export concentration, the larger the commodity price elasticity of real exchange rate, consistent with the empirical finding by Bodart et al. (2012). From the market share interaction term, we obtain a positive coefficient indicating that the monopoly pricing power of a commodity-exporting country tends to make a transmission of *RCP* shocks to the *REER* larger. Estimation results including all conditional variables are reported in column (5) just as a robustness check where we find the similar results to the earlier specifications.

INSERT TABLE 2 HERE

The interaction variable model tends to increase the likelihood of the multicollinearity. 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 interaction terms are jointly significant. So, for example, the null hypothesis of a F-test for the specification (2) in Table 2 is H_0: $\beta_{RCP\times TO} = \beta_{RCP\times FO} = 0$ and the alternative is H_1: at least one $\beta \neq 0$. For the all specifications in the DOLS(1,1) estimation in Table 2, 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.

4.1.2 Non-oil commodity vs. oil exporters

Although there are common features between the price of oil and the price of non-oil commodities, authors in the commodity currency literature often investigate two groups of countries separately, reflecting a general recognition of distinctive movements of oil prices.[21] The price of oil is very sensitive to changes in global business cycle as oil is the most widely used industrial input. At the same time, oil prices are under the influence of an oil cartel such as OPEC (Organization of the Petroleum Exporting Countries). We thus attempt to separate oil countries from the non-oil commodity exporters and look at if there exist any noticeable

[21] See Coudert et al. (2008) for an extensive literature review.

differences in real exchange rate responses to commodity/oil price changes. Countries included in oil exporters are the ones whose oil share in aggregate commodity exports is greater than 50% on average over the sample period 1980-2010. They are Algeria, Kuwait, Libya, Mexico, Nigeria, Norway, Oman, Saudi Arabia, Syria, Tunisia, United Arab Emirates, and Venezuela.

INSERT TABLE 3 HERE

As shown in Table 3, a commodity currency phenomenon is much stronger in non-oil commodity exporters than oil exporters in terms of both the magnitude of the *RCP* elasticity and statistical significance. Turning to a role of structural/policy factors, inflation targeting appears effective for non-oil commodity exporters in that it dampens a transmission of a commodity price shock to the exchange rate, which is not a case for oil countries. In fact, amongst all oil exporters in our sample, only Mexico and Norway adopted inflation targeting in 2001. On the other hand, trade openness and international reserves play a much bigger role for oil exporters in lowering the commodity price elasticity than the non-oil commodity exporters.

4.2 Commodity currency in the short-run

In this subsection, we investigate the workings of commodity currencies in the short-run. Table 4 shows the short-run commodity currencies across different country groups based on the error correction model (ECM). δ_0, θ_2 and λ in equation (22) are the parameters of our interest. From the estimation results using the full sample (column (1)), the error-correction term (*EC*) has an expected sign at the 1% level of statistical significance, verifying the presence of long-run cointegration. However, we note that the estimated quarterly adjustments towards the long-run equilibrium level of *REER* seem very slow, in line with the PPP puzzle argument in the literature. Furthermore, in contract to Chen and Rogoff (2003) that finds strong short-run commodity currencies in Australia and New Zealand, we find weak evidence of short-run *REER-RCP* relationship in our full sample with the poor goodness-of-fit measure.

INSERT TABLE 4 HERE

We extend our approach to a sub-sample, displayed in columns (2) and (3), and find only non-oil commodity countries to have a significant short-run *REER-RCP* relationship. We thus present the effect of interaction terms using only the non-oil country sample. As shown in columns (4)-(6) in Table 4, the majority of structural and policy factors are not effective in the short-run. Two factors worth mentioning in the short-run are financial openness and exchange rate regime. With the world financial market integration and development of financial instruments, the effect of the cross-border capital flows on commodity currencies is prominent even in the short-run. In addition, fixed exchange rate regime, which tends to amplify commodity price shocks in the long-run, seems to achieve its original policy objective of stabilization in the presence of nominal rigidity.

Overall, the magnitude of the short-run commodity price elasticity, even when significant, is much smaller than the one in the long-run. One plausible explanation for this weak real exchange rate response based on our theory is that all factors are indeed sector specific in the short-run, making the factor price adjustment nearly impossible.

4.3 Robustness checks

In this subsection, we test the robustness of our empirical results. First, we consider the potential structural shifts in the long-run cointegrating relationship between *REER* and *RCP*. Abrupt changes in real effective exchange rate would blur the cointegrating relationship between *REER* and *RCP* and need to be properly controlled. Gregory-Hansen (1996) proposes 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. Formally, the Gregory-Hansen (1996) test is based on the following model:

$$REER_t = \beta_0 + \beta_1 \vartheta_{t\pi} + \beta_2 RCP_t + \varepsilon_t \tag{23}$$

where β_1 is the coefficient of the dummy variable $\vartheta_{t\pi}$ that models a structural change as follows:

$$\vartheta_{t\pi} = \begin{cases} 0 & \text{if } t \le [N\pi] \\ 1 & \text{if } t > [N\pi] \end{cases} \tag{24}$$

where N is the sample size, $\pi \in (0,1)$ is a fraction parameter that determines a timing of the level shift, and [ ] denotes integer part. Following Gregory and Hansen (1996), we choose 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 the lag length chosen based on AIC (Akaike Information Criterion). The null hypothesis of the test is no cointegration against the alternative of cointegration with a single shift at an unknown point in time. Columns (8) and (9) 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 consistent with the country specific macroeconomic events such as hyperinflation, exchange rate crisis, and nominal exchange rate adjustment program including a remarkable devaluation of the CFA franc by 50 percent in early 1994. The regression results presented in Table 5 include dummy variables controlling for country-specific structural shift dates. Even after controlling for regime shifts in the cointegrating relationship, the DOLS(1,1) regression results in Table 5 are sufficiently close to the main results in Table 2 and we conclude that our results are robust to structural shift consideration.

INSERT TABLE 5 HERE

Next, as a robustness check for the short-run commodity currency estimation, we run an ECM regression after accounting for structural shifts. Column (1) in Table 6 shows that controlling for structural shifts does not make a much difference in the short-run commodity currency relation from the results in column (1) in Table 4. The commodity price coefficient remains insignificant while the error correction term is significant with an expected sign. We then look at the robustness of short-run empirical results for a group of non-oil commodity countries controlling for structural shift dates and they are found robust as shown in columns (2)-(5) in Table 6.[22]

[22] Additionally, we find that our long-run DOLS estimation results are robust to longer leads and lags of cointegrated *RCP* variable, validating our choice of lead and lag in a DOLS specification. For our short-run ECM estimation results, higher-order lag terms of $\Delta REER$ and ΔRCP upto $t-4$ are also considered but the main results remain not

INSERT TABLE 6 HERE

5 Concluding remarks

In this paper, we demonstrate that a commodity currency phenomenon can be generalized in a large group of developing countries. This is a significant expansion from the existing literature which often focuses only on a small set of advanced economies. Despite of the strong exchange rate-commodity price connection in general, a large degree of heterogeneity in the real exchange rate response to a commodity price change is observed at an individual country level. This paper investigates the source of heterogeneity from various perspectives.

First of all, our empirical analysis based on a non-stationary panel data set finds that the long-run response of the real exchange rate to a commodity boom, largely in accordance with our structural model prediction, is smaller when a country exhibits any of the following economic characteristics: i) capital control, ii) open trade, iii) flexible nominal exchange rate, iv) massive international reserves, v) low commodity export dependency, and vi) small share of global commodity supply. In addition, in contrast to previous studies based on the currencies of a small set of developed countries, a commodity price-exchange rate relation is found much weaker in the short-run than in the long-run. We also find the weaker relation for a group of oil-exporting countries than the non-oil commodity counterparts.

Facing a rising commodity price trend mainly driven by the strong global demand during the past decade, commodity-dependent economies have recently been exposed to a real appreciation pressure. This is likely to induce high volatility in aggregate output and the price level, consequently incurring high macroeconomic adjustment costs. Therefore, given concerns for the Dutch disease or resource curse that operate through the real exchange rate, our findings in this paper are of particular relevance for monetary policy-making and for globalization

sensitive to the inclusion of additional lags. Moreover, we try including structural/policy factors in the DOLS(1,1) regression as additional main effects. In such an exercise, we have to interpret coefficients of conditional factors and of the interaction terms together to fully understand the role of those factors. Results for these exercises are available in Online Appendix at https://sites.google.com/site/leedwec.

strategy in commodity-exporting developing economies in order for them to effectively manage costly commodity price shocks.

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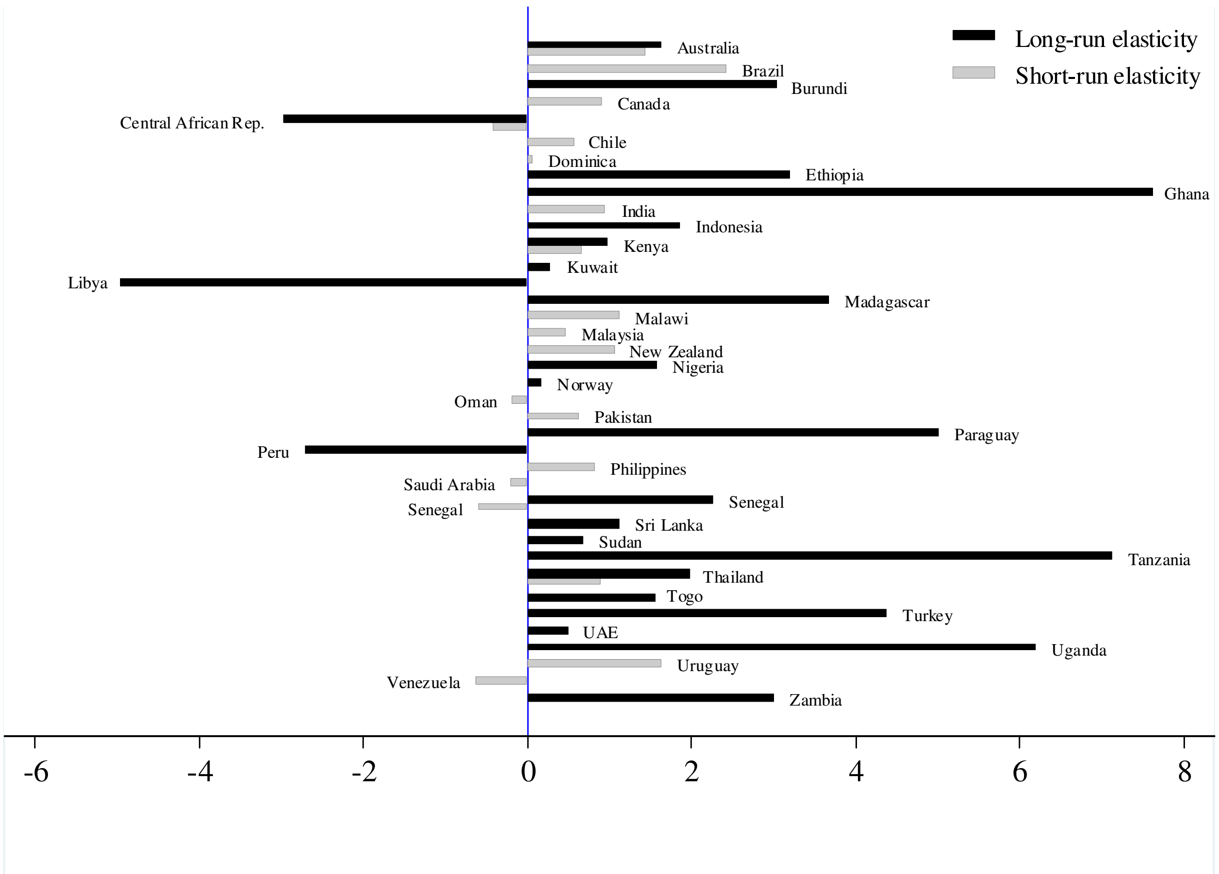
Figure 1. Distribution of commodity price elasticity of the real exchange rate across countries[23]

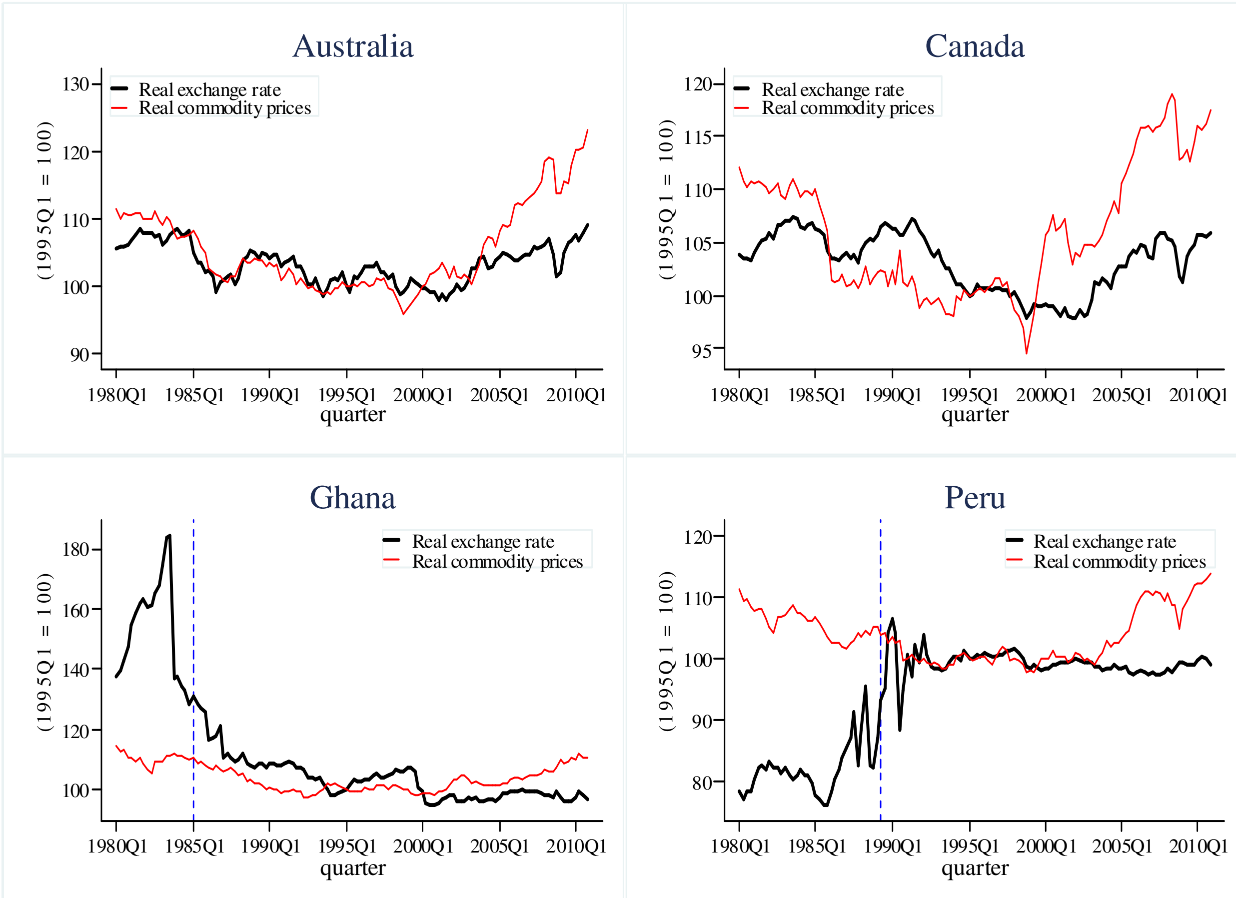
$$\text{DOLS(1,1):}\; REER_t = \alpha_0 + \beta_1 RCP_t + \sum_{j=-1}^{1} \Delta RCP_{t+j}\gamma_j + u_t$$

$$\text{First differencing:}\; \Delta REER_t = \alpha_1 + \delta_1 \Delta RCP_t + \varepsilon_t$$



[23] We estimate the long-run elasticity by the dynamic OLS and short-run elasticity by first differences based on the time series test results reported in Table A1 in Appendix. According to the country-by-country analysis, commodity prices and the real exchange rates are non-stationary but cointegrated for the majority of countries in our sample.

Figure 2. Time series plots of real effective exchange rate and real commodity prices[24]



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

Table 1: Cross-sectional dependence, panel unit-root and cointegration tests

a. Cross-sectional dependence test

Specification	CD test statistic	Average absolute value of the off-diagonal elements
(1)	121.22***	0.445
(2)	32.721***	0.139

b. Panel unit-root tests

Method	*REER*		*RCP*	
	Levels	1st differences	Levels	1st differences
LLC test				
t^*-statistic	3.520	-52.872***	8.615	-40.329***
	(0.999)	(<.01)	(1.00)	(<.01)
IPS test				
W-statistic	2.395	-48.216***	9.207	-51.619***
	(0.992)	(<.01)	(1.00)	(<.01)
CIPS test				
Z[t-bar] statistic	0.992	-7.447***	2.634	-11.718***
	(0.839)	(<.01)	(0.996)	(<.01)

c. Panel cointegration tests

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: In panel a, Pesaran (2004)'s cross-sectional dependence (CD) test statistic is based on the residuals of the regression model specifications (1) $REER_{it} = \alpha_i + \beta_1 RCP_{it} + u_{it}$ and (2) $\Delta REER_{it} = \alpha_i + \delta_1 \Delta RCP_{it} + \varepsilon_{it}$ where $REER$ and RCP are in logarithm. In panel b, 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. In panel c, for the Kao test, an individual intercept is included only, while the individual intercept and individual trend are included for the Pedroni test. For the Westerlund test, we set the width of Bartlett-kernel window at 4 and allow for a constant but no deterministic trend in the cointegrating relationship. In all panels, the associated p-values of the test statistics are given in parentheses. *** and ** indicate the rejection of the null hypotheses (cross-sectional independence, unit-root and no cointegration for panels a, b and c, respectively) at the 1 and 5 percent significance levels. Lag lengths are automatically selected based on the modified Akaike Information Criterion (MAIC) for all panel-unit root and cointegration tests except for the Westerlund test that uses AIC.

Table 2. Long-run elasticity and interaction effects: Full sample

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)
$RCP_t \times TO_t$		-0.206***			-0.123***
		(0.028)			(0.016)
$RCP_t \times FO_t$		0.043**			0.038**
		(0.017)			(0.015)
$RCP_t \times IT_t$			-0.023		0.024
			(0.025)		(0.021)
$RCP_t \times EXR_t$			0.047***		0.077***
			(0.016)		(0.015)
$RCP_t \times RES_t$			-0.078***		-0.046**
			(0.02)		(0.019)
$RCP_t \times CEX_t$				0.063***	0.058***
				(0.02)	(0.014)
$RCP_t \times MSH_t$				0.032*	-0.0004
				(0.017)	(0.018)
F-statistic		31.26***	30.59***	6.47***	14.53***
Within R^2	0.07	0.14	0.09	0.06	0.12
# of countries	63	63	63	63	63
Observations	7349	7108	6869	5946	5550

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. F-statistic and its significance level are reported to show if interaction terms are jointly significant. ***, **, * indicate significance at the 1%, 5%, and 10%, respectively.

Table 3. Long-run elasticity and interaction effects: Non-oil commodity vs. oil exporters

Dependent variable: *$REER_t$*	Non-oil commodity exporters				Oil exporters			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
RCP_t	1.526***	1.492***	1.451***	0.931***	0.035	0.83**	0.798*	0.086
	(0.441)	(0.421)	(0.379)	(0.220)	(0.337)	(0.341)	(0.414)	(0.232)
$RCP_t \times TO_t$		-0.164***				-0.761***		
		(0.024)				(0.129)		
$RCP_t \times FO_t$		0.049***				-0.018		
		(0.016)				(0.071)		
$RCP_t \times IT_t$			-0.042**				0.045	
			(0.020)				(0.068)	
$RCP_t \times EXR_t$			0.055***				-0.045	
			(0.014)				(0.094)	
$RCP_t \times RES_t$			-0.053***				-0.509***	
			(0.016)				(0.100)	
$RCP_t \times CEX_t$				0.054***				0.085
				(0.016)				(0.096)
$RCP_t \times MSH_t$				0.049**				-0.037
				(0.020)				(0.032)
F-statistic		26.37***	29.82***	8.64***		17.71***	16.39***	2.19
Within R^2	0.12	0.17	0.13	0.09	0.03	0.29	0.18	0.03
# of countries	51	51	51	51	12	12	12	12
Observations	5927	5790	5738	4805	1422	1318	1131	1141

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. *F*-statistic and its significance level are reported to show if interaction terms are jointly significant. ***, **, * indicate significance at the 1%, 5%, and 10%, respectively.

Table 4. Short-run elasticity, dynamic adjustment and interaction effects

Dependent variable: $\Delta REER_t$						
	Full	Non-oil	Oil	Non-oil		
	(1)	(2)	(3)	(4)	(5)	(6)
ΔRCP_t	0.057	0.145***	-0.081	0.137	0.37***	0.207**
	(0.043)	(0.046)	(0.084)	(0.119)	(0.133)	(0.101)
EC_{t-1}	-0.043***	-0.052***	-0.023*	-0.054***	-0.057***	-0.059***
	(0.011)	(0.014)	(0.011)	(0.015)	(0.015)	(0.010)
$\Delta REER_{t-1}$	-0.022	-0.034	0.055	-0.039	-0.041	-0.042
	(0.029)	(0.028)	(0.075)	(0.029)	(0.028)	(0.042)
ΔRCP_{t-1}	-0.0004	0.013	-0.011	0.02	0.006	0.003
	(0.042)	(0.055)	(0.056)	(0.055)	(0.057)	(0.047)
$\Delta RCP_t \times TO_t$				-0.112		
				(0.112)		
$\Delta RCP_t \times FO_t$				0.185**		
				(0.084)		
$\Delta RCP_t \times IT_t$					0.39	
					(0.257)	
$\Delta RCP_t \times EXR_t$					-0.329***	
					(0.111)	
$\Delta RCP_t \times RES_t$					-0.02	
					(0.077)	
$\Delta RCP_t \times CEX_t$						-0.121
						(0.12)
$\Delta RCP_t \times MSH_t$						0.136
						(0.126)
F-statistic				3.04*	4.79***	2.02
Within R^2	0.03	0.04	0.02	0.04	0.04	0.04
# of countries	63	51	12	51	51	51
Observations	7400	5968	1432	5828	5776	4842

Note: Column (1) shows the ECM estimation results using the full sample. Countries included in columns (2) and (3) are non-oil commodity exporters and oil exporters, respectively. Estimation results in columns (4)-(6) include non-oil commodity exporting countries only. Country fixed effects are included in all specifications. Driscoll-Kraay standard errors are reported in parentheses. *F*-statistic and its significance level are reported to show if interaction terms are jointly significant. ***, **, * indicate significance at the 1%, 5%, and 10%, respectively.

Table 5. Robustness I: Long-run results controlling for structural shifts

Dependent variable: $REER_t$					
	(1)	(2)	(3)	(4)	(5)
RCP_t	0.93**	1.024**	1.021***	0.587**	0.559***
	(0.415)	(0.411)	(0.368)	(0.228)	(0.198)
$RCP_t \times TO_t$		-0.206***			-0.123***
		(0.029)			(0.016)
$RCP_t \times FO_t$		0.043**			0.037**
		(0.017)			(0.015)
$RCP_t \times IT_t$			-0.023		0.024
			(0.025)		(0.021)
$RCP_t \times EXR_t$			0.047***		0.077***
			(0.016)		(0.015)
$RCP_t \times RES_t$			-0.079***		-0.046**
			(0.02)		(0.019)
$RCP_t \times CEX_t$				0.063***	0.058***
				(0.02)	(0.014)
$RCP_t \times MSH_t$				0.031*	-0.001
				(0.017)	(0.018)
F-statistic		30.62***	30.61***	6.45***	14.45***
Within R^2	0.07	0.14	0.10	0.06	0.12
# of countries	63	63	63	63	63
Observations	7349	7108	6869	5946	5550

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 includes country fixed effects as well as level shift dummies to control for structural shift dates identified by the Gregory and Hansen (1996) test (Bolivia, 1985q4; Burundi, 2002q4; Cameroon, 1993q2; Central African Republic, 1993q2; Costa Rica, 1992q1; Ethiopia, 1993q1; Ghana, 1985q1; Kenya, 2000q4; Libya, 1994q1; Madagascar, 1986q4; Norway, 1992q2; Oman, 1986q1; Papua New Guinea, 1998q3; Paraguay, 1987q3; Peru, 1989q2; Saudi Arabia, 1986q1; Senegal, 1993q3; Syria, 1989q2; Togo, 1993q2; Tunisia, 1986q3; Uganda, 1990q1; Zambia, 1987q4). Driscoll-Kraay standard errors are reported in parentheses. *F*-statistic and its significance level are reported to show if interaction terms are jointly significant. ***, **, * indicate significance at the 1%, 5%, and 10%, respectively.

Table 6. Robustness II: Short-run results controlling for structural shifts

Dependent variable: $\Delta REER_t$					
	Full	Non-oil			
	(1)	(2)	(3)	(4)	(5)
ΔRCP_t	0.051	0.144***	0.136	0.371***	0.202*
	(0.041)	(0.046)	(0.118)	(0.131)	(0.103)
EC_{t-1}	-0.043***	-0.051***	-0.053***	-0.056***	-0.058***
	(0.011)	(0.014)	(0.015)	(0.015)	(0.01)
$\Delta REER_{t-1}$	-0.031	-0.046	-0.052	-0.053	-0.067
	(0.03)	(0.032)	(0.033)	(0.033)	(0.05)
ΔRCP_{t-1}	0.0001	0.015	0.024	0.009	0.009
	(0.042)	(0.056)	(0.056)	(0.058)	(0.049)
$\Delta RCP_t \times TO_t$			-0.115		
			(0.114)		
$\Delta RCP_t \times FO_t$			0.191**		
			(0.085)		
$\Delta RCP_t \times IT_t$				0.399	
				(0.256)	
$\Delta RCP_t \times EXR_t$				-0.329***	
				(0.111)	
$\Delta RCP_t \times RES_t$				-0.026	
				(0.076)	
$\Delta RCP_t \times CEX_t$					-0.116
					(0.12)
$\Delta RCP_t \times MSH_t$					0.148
					(0.126)
F-statistic			3.07*	4.82***	2.12
Within R^2	0.03	0.04	0.04	0.05	0.05
# of countries	63	51	51	51	51
Observations	7400	5968	5828	5776	4842

Note: Column (1) reports the ECM estimation results for the full sample and columns (2)-(5) show results for non-oil commodity exporting countries only. All specifications include country fixed effects as well as level shift dummies to control for structural shift dates identified by the Gregory and Hansen (1996) test (Bolivia, 1985q4; Burundi, 2002q4; Cameroon, 1993q2; Central African Republic, 1993q2; Costa Rica, 1992q1; Ethiopia, 1993q1; Ghana, 1985q1; Kenya, 2000q4; Libya, 1994q1; Madagascar, 1986q4; Norway, 1992q2; Oman, 1986q1; Papua New Guinea, 1998q3; Paraguay, 1987q3; Peru, 1989q2; Saudi Arabia, 1986q1; Senegal, 1993q3; Syria, 1989q2; Togo, 1993q2; Tunisia, 1986q3; Uganda, 1990q1; Zambia, 1987q4). Driscoll-Kraay standard errors are reported in parentheses. F-statistic and its significance level are reported to show if interaction terms are jointly significant. ***, **, * indicate significance at the 1%, 5%, and 10%, respectively.

Appendix

1 Choice of sample countries

We keep commodity-dependent developing countries whose export earnings in nonfuel primary products accounted for more than half of total export earnings for the years 1988-1992.[25] From 73 countries based on this classification, 21 countries were excluded because times series data on either the real effective exchange rate or UN COMTRADE commodity exports are not available for a sufficiently long period of time. In addition, 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. Zimbabwe was dropped as well due to the hyperinflation during the significant part of sample period (since 2002) that could distort an appropriate measure of the exchange rate. Five commodity-dependent developed countries (Australia, Canada, Iceland, Norway and New Zealand) and nine major oil exporters (Algeria, Bahrain, Kuwait, Libya, Nigeria, Oman, Saudi Arabia, United Arab Emirates, and Venezuela) were added. This procedure leaves a total of 63 commodity republics including both non-oil and oil exporters. Note that 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 Variable definitions

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

We obtain the CPI based real effective exchange rate (*REER*; base 2005 = 100), 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

[25] This is the classification originally set in the International Monetary Fund's World Economic Outlook (IMF, 1996) and adopted in Cashin et al. (2004).

commodity exports relative to the world price of manufactured goods exports. It is a common practice to measure the terms-of-trade of countries with high commodity export dependence in this way because the majority of their imports are manufactured goods that usually account for more than half of their total imports.[26] The annual commodity trade data are taken from the UN COMTRADE and the monthly world commodity price series 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-exporting countries.[27] So, for each country *i*, commodity *j*, and time *t*, a country specific index of nominal commodity price is defined as

$$\ln NCP_{it} = \sum_{j=1}^{J} W_{ij}\left(\ln P_{jt}\right), \text{ where } W_{ij} = \frac{\frac{1}{T}\sum_{t=1}^{T} X_{ij,t}}{\frac{1}{T}\sum_{t=1}^{T} CX_{it}} \tag{25}$$

where *X* is the export volume of individual commodity *j* and *CX* is the volume of the total commodity exports. The weights (*W*) remain constant over time in order to eliminate the quantity effect from the price index. This definition is similar to Cashin et al. (2004) but the difference is that we use the period-average export values of each commodity and aggregate commodity 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.

2.2 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 (2006).[28] 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*

[26] Ricci et al. (2008) alternatively use a commodity-based terms-of-trade index which is defined as the ratio of aggregate indexes of commodity exports and imports. In their real exchange rate regression estimation using a sample of 48 countries, the commodity terms-of-trade coefficient shows an expected positive sign at the 1 percent level of statistical significance.

[27] We include all traded commodities as long as their prices are available in the IMF Primary Commodity Prices and World Bank Pink Sheet. However, platinum, plywood and steel are excluded because we have no information about the corresponding SITC codes. See Table OA1 in Online Appendix at https://sites.google.com/site/leedwec for a list of commodities employed in the construction of *RCP* indices.

[28] A data set for financial openness index is from http://web.pdx.edu/~ito/Chinn-Ito_website.htm.

Restrictions (AREAER)."[29] The index runs from -1.84 to 2.48, where higher values indicate that a country is more open to cross-border capital transactions.

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 dependency in our empirical procedure. The data are collected from the World Bank WDI.

2.3 Monetary/exchange rate policy variables

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 are not an exception.[30] 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. (2010; called IRR hereafter) and use their coarse classification for a country's exchange rate regime choice.[31] This has six regimes, namely, hard peg, soft peg, managed floating, freely floating, freely falling, and dual market. 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 (total reserves excluding gold) from IMF IFS and nominal GDP from World Bank WDI to construct a *RES* (= international reserves / NGDP) variable.

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[29] 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 (2006) 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.

[30] See Table OA2 in Online Appendix for inflation targeting countries in our sample and their policy adoption dates.

[31] An updated classification is obtained from Ilzetzki's webpage at http://personal.lse.ac.uk/ilzetzki/IRRBack.htm.

2.4 Export market structure variables

Commodity Export Dependency (*CEX*). We define a country's commodity export dependency as follows: for each country *i* and time *t*,

$$CEX_{it} = \frac{\text{Total commodity exports}_{it}}{\text{Total goods exports}_{it}} \tag{26}$$

Note that a high value of *CEX* indicates a country's heavy reliance on commodity exports and a low degree of export diversification.[32]

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

$$MSH_{it} = \sum_{j=1}^{J} W_{ij,t}\left(\text{Share}_{ij,t}\right) \tag{27}$$

$$\text{where} \quad W_{ij,t} = \frac{\text{Commodity exports}_{ij,t}}{\text{Total commodity exports}_{it}};$$

$$\text{Share}_{ij,t} = \frac{\text{Commodity exports}_{ij,t}}{\text{World supply of commodity exports}_{jt}}$$

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 the transmission of a commodity price shock to the country's currency value.

[32] The data for total goods exports by each country are obtained from the UN COMTRADE.

Table A1: Commodity price elasticity estimates, unit-root and cointegration tests

Country	Elasticity estimates		DF-GLS unit-root test				Cointegration test		
			REER		*RCP*				
	DOLS	1st differencing	Trend	No Trend	Trend	No Trend	AEG $Z(t)$	G-H $Z(t)$	Shift date
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Algeria	-0.84 (0.57)	0.03 (0.33)	-1.58 (1)	0.04 (1)	-0.75 (5)	-0.71 (5)	-1.11 (1)	-4.59*	
Argentina		0.05 (0.46)	-2.09 (1)	-1.16 (1)	-1.44 (1)	-1.35 (1)	-2.69 (1)	-3.35	
Australia	1.64*** (0.24)	1.46*** (0.41)	-1.38 (2)	-1.37 (2)	0.04 (1)	-0.16 (1)	-3.08* (1)	-4.54*	
Bahrain		0.05 (0.12)	-2.35 (1)	-0.42 (11)	-1.15 (1)	-1.12 (1)	-0.58 (1)	-3.5	
Bangladesh		0.08 (0.15)	-1.36 (2)	-1.55 (2)	-1.41 (8)	0.06 (7)	-1.34 (2)	-3.19	
Bolivia	1.27 (0.87)	1.73 (1.27)	-1.61 (12)	-0.73 (12)	-0.55 (6)	-0.66 (6)	-1.79 (3)	-7.7***	1985q4
Brazil		2.44*** (0.66)	-2.09 (1)	-2.02* (1)	-0.01 (3)	-0.41 (3)	-2.48 (1)	-2.97	
Burundi	3.04*** (0.35)	0.17 (0.24)	-2.41 (1)	-0.89 (1)	-0.75 (3)	-1.01 (11)	-1.30 (1)	-4.66**	2002q4
Cameroon	-0.53 (0.32)	-0.13 (0.09)	-1.93 (1)	-1.00 (1)	-0.75 (2)	-0.88 (2)	-1.74 (1)	-4.66**	1993q2
Canada		0.92*** (0.30)	-1.64 (3)	-1.68 (3)	-0.67 (1)	-0.75 (1)	-1.58 (1)	-3.44	
Central African Rep.	-2.98** (1.14)	-0.43** (0.21)	-1.64 (1)	-0.50 (1)	-1.79 (10)	-1.31 (10)	-1.55 (2)	-5.06**	1993q2
Chile		0.58** (0.23)	-0.56 (2)	-0.26 (2)	-1.27 (1)	-1.26 (1)	-2.16 (1)	-3.93	
Colombia		0.35* (0.19)	-1.30 (1)	-0.86 (1)	-0.40 (2)	-0.70 (2)	-1.38 (1)	-3.69	
Costa Rica	0.58* (0.34)	-0.05 (0.13)	-0.42 (8)	-0.58 (8)	-0.99 (7)	-0.70 (7)	-5.77*** (4)	-5.58***	1992q1
Cote d'Ivoire		-0.13 (0.24)	-2.32 (1)	-1.30 (1)	-0.60 (1)	-0.63 (1)	-2.65 (1)	-3.8	
Dominica		0.08** (0.04)	-1.60 (2)	-1.48 (1)	-1.56 (11)	-1.43 (11)	-1.17 (1)	-2.85	
Ethiopia	3.22*** (0.57)	-0.57 (0.61)	-1.95 (1)	-0.50 (1)	-1.25 (5)	-1.09 (11)	-1.58 (1)	-5.24***	1993q1
Ghana	7.63*** (1.70)	0.61 (0.76)	-3.71*** (1)	-1.38 (1)	-0.42 (9)	-0.60 (9)	-1.36 (1)	-4.84**	1985q1
Guatemala		0.21 (0.25)	-1.27 (1)	-1.15 (1)	-0.62 (5)	-0.72 (5)	-2.49 (1)	-3.34	
Honduras		-0.003 (0.26)	-1.39 (4)	-1.34 (4)	-1.01 (7)	-0.62 (7)	-2.51 (2)	-4.11	
Iceland		0.06 (0.43)	-2.19 (6)	-0.34 (6)	-1.46 (1)	-0.71 (1)	-1.80 (1)	-2.66	
India		0.97*** (0.34)	-0.91 (4)	-0.59 (4)	-0.34 (4)	-0.66 (4)	-0.42 (1)	-4.08	
Indonesia	1.87** (0.78)	0.54 (0.57)	-1.74 (4)	-0.89 (1)	-0.58 (1)	-0.74 (1)	-1.89 (1)	-4.46*	
Kenya	1.00*** (0.26)	0.67** (0.28)	-0.69 (6)	-0.68 (6)	-1.46 (5)	-0.86 (5)	-1.59 (1)	-4.98**	2000q4
Kuwait	0.29*** (0.09)	0.19 (0.17)	-2.17 (4)	-1.44 (1)	-0.76 (5)	-0.77 (5)	-2.19 (1)	-4.45*	
Libya	-4.96*** (0.24)	-0.33 (0.33)	-1.44 (1)	-0.28 (5)	-0.77 (5)	-0.78 (5)	-3.04 (1)	-5.82***	1994q1
Madagascar	3.69*** (0.56)	-0.58 (0.36)	-1.95 (1)	-0.89 (1)	-1.96 (5)	-0.46 (5)	-2.78 (1)	-5.23***	1986q4
Malawi		1.14** (0.51)	-3.79*** (1)	0.01 (6)	-1.42 (1)	-1.04 (1)	-1.41 (2)	-4.06	
Malaysia		0.48** (0.19)	-2.05 (2)	-0.35 (1)	-0.67 (1)	-0.84 (1)	-1.30 (1)	-4.05	
Mali		-0.59 (0.44)	-0.81 (2)	0.44 (3)	0.21 (5)	-0.49 (1)	-0.11 (3)	-3.48	
Mauritania		0.14 (0.14)	-2.10 (1)	0.18 (1)	-0.07 (4)	0.10 (4)	-1.28 (1)	-2.84	
Mauritius		0.04 (0.06)	-1.92 (2)	-0.70 (1)	-1.79 (1)	-1.51 (1)	-1.93 (1)	-4.33	
Mexico		0.30 (0.20)	-2.52 (1)	-2.32** (1)	-0.61 (5)	-0.72 (5)	-2.69 (1)	-3.1	
Morocco		0.02 (0.08)	-0.79 (5)	0.44 (5)	-0.98 (12)	-0.93 (12)	-3.03 (1)	-4.22	
Mozambique		0.16 (0.39)	-1.88 (2)	-1.73 (2)	-2.25 (1)	-1.06 (1)	-2.08 (3)	-3.53	

Table A1 (continued)

Country	Elasticity estimates		DF-GLS unit-root test				Cointegration tests		
	DOLS (1,1)	1st differencing	*REER*		*RCP*		AEG	G-H $Z(t)$	Shift date
			Trend	No Trend	Trend	No Trend			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
New Zealand		1.07*** (0.38)	-2.44 (4)	-1.93* (4)	-1.02 (7)	-0.64 (7)	-2.49 (1)	-3.33	
Niger		0.02 (0.12)	-0.91 (1)	0.34 (1)	-0.48 (1)	-0.72 (1)	-1.16 (1)	-3.69	
Nigeria	1.58*** (0.54)	-0.35 (0.29)	-1.57 (1)	-0.89 (1)	-0.75 (5)	-0.77 (5)	-1.67 (1)	-4.35*	
Norway	0.19*** (0.05)	0.28 (0.18)	-1.98 (5)	-1.82* (5)	-0.69 (5)	-0.73 (5)	-3.13*(1)	-4.71**	1992q2
Oman	-0.16 (0.30)	-0.19** (0.08)	-1.95 (2)	-0.21 (7)	-0.74 (5)	-0.76 (5)	-1.29 (1)	-4.94**	1986q1
Pakistan		0.64*** (0.18)	-0.67 (2)	0.39 (1)	-0.73 (6)	-0.68 (6)	-1.07 (1)	-3.18	
Papua New Guinea	0.51 (0.34)	-0.24 (0.26)	-0.78 (8)	-0.89 (1)	-0.48 (1)	-0.78 (1)	-1.73 (1)	-4.92**	1998q3
Paraguay	5.02*** (0.52)	-0.08 (0.38)	-1.08 (3)	-0.48 (3)	-1.83 (2)	-1.44 (2)	-3.78** (1)	-4.87**	1987q3
Peru	-2.72*** (0.81)	-0.70 (0.80)	-1.29 (11)	-0.15 (11)	-0.46 (2)	-0.78 (2)	-1.81 (1)	-7.03***	1989q2
Philippines		0.84*** (0.28)	-1.14 (6)	-0.55 (6)	-0.41 (10)	-0.69 (10)	-1.92 (1)	-3.84	
Saudi Arabia	0.06 (0.34)	-0.20** (0.08)	-1.21 (1)	0.12 (3)	-0.76 (5)	-0.77 (5)	-1.54 (1)	-5.48***	1986q1
Senegal	2.28*** (0.48)	-0.60** (0.29)	-1.86 (1)	-0.12 (1)	-0.08 (6)	-0.54 (6)	-0.96 (1)	-5.45***	1993q3
South Africa		0.44 (0.64)	-2.28 (2)	-1.19 (3)	-0.21 (2)	-0.29 (2)	-2.08 (1)	-3.54	
Sri Lanka	1.14*** (0.20)	0.17 (0.14)	-1.52 (1)	-1.02 (1)	-1.40 (12)	-1.01 (12)	-2.01 (1)	-4.6*	
St. Vincent Gr	0.33 (0.40)	0.11 (0.07)	-1.99 (4)	-1.34 (4)	-0.85 (11)	-0.83 (11)	-0.93 (4)	-4.44*	
Sudan	0.69*** (0.06)	0.03 (0.07)	-2.13 (4)	0.42 (5)	-0.70 (5)	-0.75 (5)	-3.37* (1)	-4.05	
Suriname		0.48 (0.41)	-0.88 (1)	-0.83 (1)	-0.91 (1)	-0.92 (1)	-1.86 (1)	-2.5	
Syria	-0.72 (0.60)	0.004 (0.27)	-1.64 (1)	-1.19 (1)	-0.73 (5)	-0.78 (5)	-1.25 (1)	-4.78**	1989q2
Tanzania	7.14*** (1.18)	1.39 (1.05)	-2.00 (1)	-1.00 (1)	0.13 (4)	-0.64 (4)	-0.89 (1)	-4.51*	
Thailand	2.01*** (0.39)	0.98*** (0.35)	-1.57 (2)	-0.63 (2)	-0.32 (1)	-0.74 (1)	-1.55 (1)	-4.39*	
Togo	1.58*** (0.56)	-0.37 (0.32)	-1.58 (4)	-0.36 (3)	-0.33 (6)	-0.76 (10)	-1.98 (2)	-5.31***	1993q2
Tunisia	-0.09 (0.41)	0.05 (0.07)	-1.18 (1)	0.07 (5)	-0.79 (5)	-0.81 (5)	-1.47 (1)	-6.25***	1986q3
Turkey	4.40*** (0.84)	0.64 (0.84)	-2.47 (5)	0.17 (4)	-0.37 (1)	-0.69 (1)	-3.54** (1)	-4.32	
Uganda	6.21*** (0.84)	0.15 (1.25)	-0.96 (6)	-0.07 (10)	-0.90 (5)	-0.99 (11)	-3.85** (1)	-4.85**	1990q1
United Arab Emirates	0.50*** (0.07)	-0.04 (0.09)	-2.25 (1)	-1.51 (1)	-0.62 (5)	-0.70 (5)	-3.51** (1)	-3.27	
Uruguay		1.65*** (0.41)	-1.06 (2)	-1.13 (2)	-0.36 (5)	-0.55 (5)	-1.89 (1)	-3.12	
Venezuela, RB		-0.63** (0.26)	-1.46 (1)	-1.37 (1)	-0.74 (5)	-0.77 (5)	-3.01 (1)	-3.83	
Zambia	3.02*** (0.32)	0.54 (0.39)	-2.43 (3)	-0.08 (3)	-1.41 (1)	-1.36 (1)	-3.60** (1)	-6.07***	1987q4

Note: Columns (1) and (2) present commodity price elasticity estimates with Newey-West HAC standard errors in brackets. Structural shift dates, which are reported in column (9), are controlled in the estimation procedure. Columns (3)-(6) report test statistics of DF-GLS unit-root test (Elliot et al., 1996) for the real effective exchange rate and real commodity prices with and without a deterministic trend term. The lag length is automatically chosen due to the minimum of the modified Akaike information criterion (MAIC) and presented in parentheses. Column (7) presents the Augmented Engel-Granger (AEG) cointegration test statistic and its level of significance (based on the critical values from MacKinnon (1990, 2010)) with the number of optimal lags chosen by the Schwarz Bayesian information criterion (SBIC) reported in parentheses. Columns (8) and (9) report the Gregory and Hansen (1996) test statistics and associated structural shift dates. For all columns, ***, **, * indicate statistical significance at the 1%, 5%, and 10% level, respectively.

Table A2. Descriptive statistics

Variable	Obs	Mean	Std. Dev.	Min	Max	Median
REER	7526	4.774	0.391	3.362	8.615	4.684
RCP	7812	1.212	0.448	0.302	3.992	1.149
TO	7552	0.652	0.343	0.063	2.511	0.586
FO	7748	-0.179	1.436	-1.864	2.439	-1.169
IT	7812	0.094	0.292	0	1	0
EXR	7326	2.401	1.251	1	6	2
RES	7584	0.112	0.118	0	1.583	0.087
CEX	6168	0.490	0.258	0.01	0.997	0.505
MSH	6328	0.079	0.081	0	0.630	0.049

Table A3. Primary exporting commodities and their share in aggregate commodity exports

Country	Primary commodities					Share in commodity exports				
	1	2	3	4	5	1	2	3	4	5
Algeria	Crude oil	Natural gas				0.59	0.41			
Argentina	Soy meal	Wheat	Maize	Soybeans	Crude oil	0.18	0.12	0.10	0.10	0.09
Australia	Coal	Iron	Beef	Gold	Wheat	0.21	0.11	0.09	0.09	0.08
Bahrain	Aluminum	Natural gas	Crude oil	Iron	Urea	0.74	0.40	0.27	0.17	0.14
Bangladesh	Shrimp	Tea	Urea	Fish	Beef	0.69	0.27	0.12	0.08	0.05
Bolivia	Natural gas	Zinc	Tin	Soy meal	Gold	0.40	0.15	0.10	0.07	0.07
Brazil	Iron	Coffee	Soy meal	Natural gas	Soybeans	0.18	0.13	0.10	0.10	0.09
Burundi	Coffee	Gold	Tea	Sugar	Hides	0.56	0.39	0.04	0.02	0.01
Cameroon	Crude oil	Cocoa	Coffee	Aluminum	Hard sawnwood	0.48	0.13	0.10	0.07	0.06
Canada	Crude oil	Natural gas	Soft sawnwood	Woodpulp	Wheat	0.17	0.16	0.12	0.10	0.07
Central African Rep	Hard logs	Cotton	Hard sawnwood	Coffee	Soft logs	0.38	0.32	0.17	0.15	0.05
Chile	Copper	Natural gas	Woodpulp	Fish	Fishmeal	0.67	0.11	0.07	0.07	0.06
Colombia	Coffee	Crude oil	Coal	Bananas	Gold	0.39	0.37	0.14	0.07	0.03
Costa Rica	Bananas	Coffee	Fish	Beef	Natural gas	0.48	0.34	0.06	0.06	0.05
Cote d'Ivoire	Cocoa	Coffee	Crude oil	Hard sawnwood	Rubber	0.49	0.14	0.11	0.07	0.05
Dominica	Bananas	Oranges	Coconut oil	Soy oil		0.94	0.04	0.01	0.01	
Ethiopia	Coffee	Hides	Gold	Sugar	Beef	0.79	0.11	0.10	0.02	0.01
Ghana	Gold	Cocoa	Natural gas	Hard sawnwood	Aluminum	0.45	0.38	0.28	0.07	0.05
Guatemala	Coffee	Sugar	Bananas	Natural gas	Crude oil	0.41	0.19	0.15	0.14	0.08
Honduras	Coffee	Bananas	Shrimp	Palm oil	Sugar	0.40	0.30	0.09	0.05	0.03
Iceland	Fish	Aluminum	Fishmeal	Shrimp	Beef	0.60	0.26	0.09	0.04	0.01
India	Iron	Rice	Shrimp	Tea	Crude oil	0.18	0.13	0.12	0.12	0.09
Indonesia	Crude oil	Natural gas	Rubber	Copper	Coal	0.35	0.22	0.07	0.06	0.06
Kenya	Tea	Coffee	Fish	Palm oil	Gold	0.53	0.32	0.04	0.02	0.01
Kuwait	Crude oil	Natural gas	Urea	Gold	Shrimp	0.95	0.34	0.09	0.03	0.01
Libya	Crude oil	Natural gas	Urea			0.98	0.02	0.01		
Madagascar	Shrimp	Coffee	Sugar	Cocoa	Hard sawnwood	0.52	0.39	0.06	0.04	0.02
Malawi	Tobacco	Tea	Sugar	Uranium	Coffee	0.68	0.12	0.11	0.07	0.02
Malaysia	Crude oil	Palm oil	Natural gas	Rubber	Hard logs	0.28	0.20	0.16	0.10	0.08
Mali	Gold	Cotton	Lamb	Groundnut oil		0.56	0.48	0.04	0.01	
Mauritania	Iron	Fish	Crude oil	Copper	Gold	0.63	0.27	0.16	0.15	0.13
Mauritius	Sugar	Fish	Tea	Wheat		0.92	0.05	0.01	0.01	

Table A3 (continued)

Mexico	Crude oil	Natural gas	Coffee	Silver	Copper	0.77	0.20	0.04	0.03	0.03
Morocco	Phosphate rock	Oranges	TSP	Fish	Lead	0.40	0.18	0.10	0.09	0.05
Mozambique	Aluminum	Shrimp	Sugar	Cotton	Tobacco	0.48	0.12	0.11	0.08	0.07
New Zealand	Beef	Wool (fine)	Aluminum	Fish	Wool (coarse)	0.39	0.10	0.09	0.07	0.06
Niger	Uranium	Gold	Lamb	Rice	Sugar	0.81	0.12	0.09	0.01	0.01
Nigeria	Crude oil	Natural gas	Cocoa			0.97	0.21	0.01		
Norway	Crude oil	Natural gas	Aluminum	Fish	Nickel	0.59	0.21	0.08	0.06	0.02
Oman	Crude oil	Natural gas	Copper	Fish	Urea	0.92	0.05	0.02	0.02	0.01
Pakistan	Rice	Cotton	Natural gas	Shrimp	Crude oil	0.53	0.25	0.13	0.05	0.04
Papua New Guinea	Copper	Crude oil	Gold	Coffee	Palm oil	0.39	0.22	0.17	0.11	0.06
Paraguay	Soybeans	Cotton	Beef	Soy meal	Soy oil	0.38	0.23	0.11	0.08	0.05
Peru	Copper	Gold	Fishmeal	Zinc	Lead	0.26	0.17	0.14	0.14	0.08
Philippines	Coconut oil	Copper	Bananas	Shrimp	Sugar	0.26	0.23	0.12	0.07	0.07
Saudi Arabia	Crude oil	Natural gas				0.96	0.03			
Senegal	Fish	Groundnut oil	Phosphate rock	Crude oil	Cotton	0.26	0.24	0.19	0.09	0.09
South Africa	Coal	Aluminum	Iron	Woodpulp	Oranges	0.33	0.17	0.15	0.05	0.04
Sri Lanka	Tea	Rubber	Fish	Shrimp	Tobacco	0.73	0.13	0.03	0.03	0.02
St. Vincent Gr	Bananas	Wheat	Rice	Fish		0.55	0.26	0.18	0.01	
Sudan	Crude oil	Cotton	Gold	Lamb	Beef	0.46	0.23	0.21	0.21	0.06
Suriname	Rice	Nickel	Aluminum	Silver	Soy oil	0.55	0.29	0.19	0.17	0.14
Syria	Crude oil	Cotton	Lamb	Phosphate rock	Wheat	0.82	0.08	0.04	0.02	0.02
Tanzania	Gold	Fish	Coffee	Tobacco	Cotton	0.40	0.14	0.13	0.10	0.07
Thailand	Rice	Rubber	Shrimp	Sugar	Crude oil	0.25	0.23	0.14	0.10	0.04
Togo	Phosphate rock	Cotton	Cocoa	Coffee	Gold	0.46	0.26	0.09	0.08	0.07
Tunisia	Crude oil	Olive oil	TSP	Phosphate rock	Shrimp	0.59	0.17	0.11	0.03	0.03
Turkey	Tobacco	Aluminum	Wheat	Lamb	Gold	0.22	0.11	0.11	0.09	0.08
Uganda	Coffee	Fish	Gold	Tea	Tobacco	0.49	0.16	0.09	0.07	0.07
United Arab Em	Crude oil	Aluminum	Natural gas	Gold	Rice	0.62	0.37	0.16	0.08	0.04
Uruguay	Beef	Rice	Fish	Wool (coarse)	Soybeans	0.41	0.20	0.11	0.07	0.05
Venezuela, RB	Crude oil	Natural gas	Aluminum	Iron	Coal	0.90	0.08	0.06	0.01	0.01
Zambia	Copper	Sugar	Cotton	Tobacco	Maize	0.86	0.04	0.03	0.03	0.01

Note: Reported are top five major commodities exported by each country between 1980 and 2010. Period-average shares of each commodity in total commodity exports greater than or equal to 0.01 (1%) are included only. We admit that major commodities listed for South Africa may not well represent its actual export basket due to underreporting of gold exports during the sample period. Calculations in this table are based solely on the data available from the UN COMTRADE.