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Market power, inflation targeting, and commodity currencies

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Abstract

The 'commodity currency' literature highlights the robust exchange rate response to fluctuations in global commodity prices that occurs for major commodity exporters. The magnitude of the response, however, varies widely across countries and over time horizons. This paper examines the real exchange rates of 51 commodity exporters over the period from 1980 to 2010, and finds that in the long-run, a higher degree of market power in the world commodity trade can reduce the exchange rate response, while in the short-run, an inflation targeting regime can amplify it. These differential impacts across countries and horizons are of particular relevance for monetary policy-making and for trade strategy in commodity-abundant economies.

Keywords: Commodity price; Inflation targeting; Market power; Panel cointegration; Real exchange rate

JEL classification: C32; F31; F41; O13

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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 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.

While the literature highlights a generally robust exchange rate response to commodity price movements, especially for commodity exporters with a floating nominal exchange rate, less attention has been paid to the wide range of response magnitudes and the reasons behind it.[2] This paper seeks to understand this variation by focusing on the relative market power of commodity exporters and their choice of inflation-targeting monetary policy, both of which have characterized a number of resource-rich economies over the past several decades.

[1] Currencies that respond significantly to changes in the world prices of these countries' primary commodity exports are called 'commodity currencies'. See Amano and van Norden (1995), Chen and Rogoff (2003, 2012), and Cashin et al. (2004) for empirical exploration covering a range of developed and developing countries.

[2] Recent papers that explore this issue include Coudert et al. (2011), Bodart et al. (2012, 2015), and Chen and Lee (2014). Our paper builds upon this prior literature.

As a preliminary illustration, Fig. 1 considers a group of non-oil commodity-exporting countries and shows their domestic currency responses to movements in the world price of their primary commodity exports.[3] Regressing country-by-country the real effective exchange rate ($REER$) on the country-specific real commodity price index (RCP), we find 27 countries out of 51 to have a statistically significant commodity price coefficient at the 10% level.[4] In particular, the long-run elasticity estimates range from 0.83 to 10.93 with a standard deviation of 2.98. The lighter bars, capturing estimates of the short-run elasticity, are generally smaller in magnitude but also exhibit a wide range of values.

INSERT FIGURE 1 HERE

What may account for this heterogeneity? There is scant literature that provides answers to this question. Three notable exceptions are worth mentioning here. Aizenman et al. (2012) emphasize Latin American central banks' active management of international reserves to insulate their short-term real exchange rates from commodity terms-of-trade (CTOT) shocks. Focusing on the long-run cointegrating relation, Bodart et al. (2012) document that the larger the share a country's leading commodity is in its total export, the stronger its real exchange rate responds to changes in the corresponding commodity export price. In their more recent work, Bodart et al. (2015) present a small open economy model with tradable and non-tradable sectors, and propose

[3] This paper focuses on commodity exporters. In commodity importing economies, imported commodities are typically used as intermediate inputs to produce final output. 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.

[4] In this paper, an increase in $REER$ indicates a real appreciation of the domestic currency relative to its trading partners. RCP is defined as the world nominal prices of a country's major commodity exports, deflated by the price index of manufactured exports of all industrial economies. See section 3.1 for more details.

that the main type of commodity export of the country, its degree of trade and financial openness, its extent of export diversification, and its exchange rate regime may all be important in shaping its real exchange rate-commodity price relationship. Their empirical evidence provides support for some of these theoretical predictions, and demonstrates that the long-run commodity-currency relation seems to be weaker for developing countries that are under a more flexible exchange rate regime, or are more open to international capital or goods trade.

This paper extends the existing literature by introducing two additional factors that may affect the commodity price elasticity of a country's real exchange rate (hereinafter 'commodity price elasticity' for short). The first is a country's market power (MP) in the global commodity export market. As discussed in Chen and Rogoff (2003) in the context of Canada, Australia, and New Zealand, countries that rely heavily on primary commodity exports have a key component of their terms-of-trade being readily observable, i.e. the world price of these commodity products in major global exchanges. For a small open economy, these terms-of-trade movements can be taken as exogenous, driven mostly by global supply and demand, rather than by internal domestic conditions. We re-examine this pure price-taker assumption and re-evaluate its applicability to a wider set of commodity-exporting countries, motivated by the following observations.

First, while none of the countries in our sample is likely to have significant market power in the *aggregate* world commodity market, some do appear to hold dominant or substantial shares in the global markets of their *specific* export products. As shown in Table 1, Australia and wool, Philippines and coconut oil, Malaysia and palm oil, are three examples where the country represents more than half of the total world productions for the respective primary commodity product. Looking from the perspective of individual commodity markets (such as iron, palm kernel oil, soybean meal, uranium, among others), one may also argue that they can be more appropriately characterized as duopolistic or oligopolistic rather than perfectly competitive, as they have notably few large and dominant producers. We thus view that contrary to the common assumption taken in the literature, some of the world's commodity-exporting countries may not face a perfectly flat demand curve; rather, they may enjoy some degree of market power in one or sometimes several commodity sectors. In practice, we recognize that significant market power implies much more complicated strategic pricing behavior, with various game-theoretic and geo-political considerations coming into play. We do not purport to capture these complications in this paper (the main reason why our sample excludes OPEC countries). Instead, we construct a market share-based measure, capturing the observations above, to proxy pricing power and see if it affects the commodity price elasticity in the data.

INSERT TABLE 1 HERE

The second factor we bring to the literature is whether the commodity exporter has adopted a monetary policy with an inflation target (IT). As argued by Gerlach (1999), high commodity concentration is one of the factors that motivate the adoption of an IT scheme. Countries that specialize in primary commodities are subject to frequent and sizeable external disturbances as global commodity prices are highly volatile, which in turn put pressure on their real exchange rates over time. During the last couple of decades, these external pressures have caused a number of commodity-producing countries to abandon their currency pegs and adopt an IT regime to help address inflation concerns.[5] Our paper evaluates the impact of the IT adoption on the strength of the exchange rate-commodity price connection, a policy channel thus far has been missing in the

[5] This macroeconomic objective is achieved through greater transparency and accountability of central banks and their IT policy implementations that provide a credible anchor for inflation expectations.

commodity-currency literature. We note that our analysis on the real exchange rate response also complements the IT literature, which mostly focuses on average inflation rates, real output growth, and their volatility as outcome variables when evaluating the effect of inflation targeting in emerging and developing economies (see, for example, Gonçalves and Salles, 2008; Lin and Ye, 2009).

Aside from exploring the relevant mechanisms behind how market power and inflation targeting affect the behavior of commodity currencies, our paper also distinguishes explicitly the short-run versus the long-run determinants for the commodity price elasticity. This distinction, which is particularly important from a policy standpoint, has often been neglected in the literature.

Our empirical findings based on a panel cointegration methodology can be summarized as follows. First, for both the short-run and the long-run, our results confirm a strong and consistent exchange rate and commodity export price connection, generalizing the 'commodity currency phenomenon' to a large group of developing countries. In terms of the commodity price elasticities, we find a more muted long-run response in countries with higher degrees of market power, while inflation targeting has no significant long-run impact. On the other hand, the market power variable has little to no explanatory power for the short-run commodity price-exchange rate connection. Instead, we find that inflation targeting tends to magnify the short-run real exchange rate appreciation following a commodity price boom. These results are reasonably robust to a variety of alternative specifications, such as accounting for structural breaks, international financial crises, and possible endogeneity. In our full analyses, we also include other commodity price elasticity determinants, such as the degree of export diversification, exchange rate regimes, the level of international reserves, as well as trade and financial openness measures. The main results remain robust when we account for these other potential determinants.

While high commodity prices can bring in more revenues to their exporters, they may also bring about large inflows of hot money, inflationary pressure, and a deterioration of the international competitiveness of their other export-competing sectors due to the exchange rate appreciation. Conversely, a real depreciation resulting from a decline in commodity export prices may improve international competitiveness, but it can also raise the debt-servicing burden of emerging sovereigns and their domestic firms, whose external loan contracts are often denominated in foreign currencies. For these reasons, keeping a stable real exchange rate may be in the best interest of policy makers in commodity-exporting developing economies. This paper provides a positive analysis on the factors that affect the magnitude of the real exchange rate responses, in both short-term and long-term, with the hope that the findings can help inform normative policy analysis aiming to stabilize the economy and to minimize the cost of commodity price shocks.

The remainder of this paper is organized as follows. Section 2 offers a theoretical discussion about how a country's market power and inflation targeting policy might influence the commodity price elasticity. Section 3 explains the data and empirical methodologies employed in the paper. Section 4 presents the estimation results and their robustness. Section 5 concludes.

2. Theoretical considerations

2.1. Market power

The flexible-price small open economy models in Chen and Rogoff (2003), Cashin et al. (2004), and Bodart et al. (2015) show that a rise in the world price of a country's primary commodity export appreciates its real exchange rate in the long-run. This occurs because under the assumption of free labor mobility, higher labor demand from the commodity-exporting sector raises the overall wage rate, which in turn bids up the prices of non-traded goods, leading to a real exchange rate appreciation.[6] One common assumption in these models is that the commodity-exporting country is small and therefore cannot affect the world price of its exported goods. This pure price-taker assumption may not hold precisely in practice, particularly for countries that occupy a relatively large market share in the global trade of certain commodities.

How might an exporter's market power affect the commodity price elasticity of its real exchange rate? Before attempting to answer, we first note that each of the 58 commodity markets in our data has its own unique characteristics and market structure. Differences in their production technologies, storage costs, distributional patterns, entry barriers, degrees of product substitutability and so on make them closed to impossible to model collectively.[7] We therefore adopt a broad level approach to examine the role of 'market power', aiming to capture two general concepts.

To illustrate concepts, we first consider the aggregate market for a particular commodity product and its supply-demand curves. We note that the less elastic the demand is for this product, the less the exporter's exchange rate would respond to commodity price changes. A simple two-country theoretical model is presented in the Appendix to demonstrate the mechanism, but the intuition is straightforward. An inelastic demand implies little change in the quantity demanded in response to a price increase. Everything else equal and following the logic in the opening paragraph, any associated production change in the commodity-exporting country would also be smaller, so would its domestic labor demand and wage adjustments, thereby resulting in a more muted real exchange rate response. In other words, relative to a situation where the supplier country faces a

[6] These models are a straightforward extension of the canonical small open economy setup in, for example, De Gregorio and Wolf (1994) and Obstfeld and Rogoff (1996, ch.4).

[7] See Online Appendix Table 13 for the full range of products.

flat demand (i.e. a price-taker), facing a steeper demand is one way we aim to capture the concept of a country having some ‘market power.’

This aggregate view above captures no market structures and interactions on the supply side, where the producers may be playing oligopolistic strategies, segregating markets, or otherwise equilibrating the overall market supply and demand. Here again, we do not attempt to formalize an exact mechanism due to vast range of possibilities; instead, we argue that for a given product, a country that holds a significant market share is likely to have a lower commodity price elasticity, implicitly assuming that its higher market share reflects some degree of monopoly power. The intuition relies on the standard result that market equilibrium under a monopoly or monopolistic competition is determined by marginal cost and marginal revenue, and for the same increase in demand, the marginal revenue would increase less, so output would increase less and price would increase more under a monopoly setting than for competitive firms. Put it differently, for the same price increase observed, the quantity adjustment would be smaller for an exporter with more monopoly power than one who faces a flatter demand curve or is a pure price taker. We can then, again, follow the same logic to show that real exchange rate response would be more muted with more market power.

Based on these two broad perspectives, we construct a measure (described in section 3.1) to capture these concepts of ‘market power’, capturing both differences in the elasticities of the overall market demand across all the commodity markets, as well as in a country’s relative market shares within these different markets.

2.2. Inflation targeting

Is inflation targeting a preferred approach in commodity-exporting countries subject to volatile terms-of-trade shocks? The debate attracts significant attention, motivating the abandonment of exchange rate control and the adoption of an inflation target in New Zealand around 1990, as well as more recent proposals such as export-price anchoring (Frankel, 2003). While the precise implementation choices of IT policy (e.g. headline vs. core inflation targets) are beyond the scope of this discussion, one could hypothesize its general impact on the commodity price elasticity.

As commodity price booms tend to boost domestic economic activities, one would expect a typical inflation-targeter to raise the policy rate–appreciating its currency–to lessen the inflation pressure. To a certain extent, a real appreciation following a rise in CTOT could help avoid excess capital inflows and overheating of the economy and slow down the inflation pressure. However, this offsetting effect is unlikely to be complete. The resulting inflation would thus require adjustments through monetary policy tightening, which could lead to a sharper real appreciation in IT countries.

Note also that in the context of the 'open-economy trilemma', policymakers cannot independently stabilize domestic prices and keep fixed exchange rates under the free capital mobility. Therefore, it would be reasonable to expect greater short-term exchange rate reactions to commodity price shocks under the IT regime. Over time, however, increased credibility of the monetary policy under IT and flexible prices may help lower the real exchange rate reaction. So, a priori, we are uncertain whether the real exchange rate reaction to commodity price changes would be in the same direction in the long-run as in the short-run when IT is in effect.

3. Data and empirical methodology

Our empirical analysis is based on an annual panel data set of 51 non-oil commodity exporters over the period from 1980 to 2010. The selection method for sample countries is described in Online Appendix B.

3.1. Variable construction

In this section, we present the variables used in our empirical procedure and their construction. Our dependent variable is the CPI-based real effective exchange rate ($REER$; base 2005 = 100), which is the average of the bilateral real exchange rates between a country and its trading partners weighted by the respective trade shares of each trading partner. The data are obtained from the International Monetary Fund's (IMF) International Financial Statistics (IFS) and Information Notice System (INS).

The regressor of interest, the real commodity price index, is defined as the world (nominal) price of a country's commodity exports relative to the world price of manufactured goods exports. This standard measure of the terms-of-trade for commodity-exporting countries is motivated by the fact that the majority of their imports are manufactured goods. The annual commodity trade data are extracted from the UN COMTRADE database, and the monthly world commodity price series from the IMF Primary Commodity Prices and the World Bank (WB) Pink Sheet.

We construct the monthly commodity price indices using 58 commodities for the 51 commodity-exporting countries.[8] We first construct a country-specific index of nominal commodity prices and then deflate it by the world price of manufactured goods exports as follows:

$$RCP_{it} = \left[\sum_{j=1}^{J} w_{ij}\left(\ln P_{jt}\right)\right] / MUV_t \qquad (1)$$

where $w_{ij} = \left(1/T \sum_{t=1}^{T} y_{ij,t}\right) / \left(1/T \sum_{t=1}^{T} Y_{it}\right)$, $y_{ij,t}$ is country *i*'s export volume of individual commodity *j* at time *t*, and Y_{it} is the volume of the total commodity exports of country *i*, P_{jt} is commodity *j*'s global price, and MUV_t is the unit value index of manufactured exports for twenty industrial economies, taken from the IMF IFS. Following Cashin et al. (2004), we keep weights, w_{ij}, constant over time to eliminate the quantity effect from the price index calculation.[9]

Market power (MP). As discussed in Section 2.1, our purpose is to capture the effect of market power from two channels: how competitive the aggregate market for each commodity product is; and what market share a country in each of these markets has. We make the implicit assumption that the greater the market share, the larger the market power a country has within that market.

Formally, we construct the market share of supplier country i in the global market of commodity *j* at time *t* as follows:

$$MSH_{ij,t} = y_{ij,t} / Y_{jt}^{g} \qquad (2)$$

[8] We include all traded commodities as long as their prices are available in the IMF Primary Commodity Prices and WB Pink Sheet. Platinum, plywood, and steel are excluded because we have no information about their corresponding SITC codes. Table 13 in Online Appendix B documents a full list of commodities employed in creating RCP indices.
[9] The only notable difference is that we use the period-average values of y (export volume of each commodity) and Y (volume of the total commodity exports) over the entire sample period 1980-2010, while the weight calculation in Cashin et al. is based on the information over a shorter period 1988-1990.

where $y_{ij,t}$ is country *i*'s export volume of commodity *j* at time *t*, and Y_{jt}^{g} is the global supply of commodity *j* exports.[10]

Next, we proxy how competitive each commodity market is by using a commodity-specific Herfindahl-Hirschman Index (HHI), a common proxy for industry concentration. This index is then used to complement the world market share measure above, to obtain our overall market power index:

$$MP_{it} = \sum_{j=1}^{J} MSH_{ij,t} \times HHI_{jt} \tag{3}$$

where $HHI_{jt} = \sum_{i=1}^{N} \left(MSH_{ij,t}\right)^2$. The market power index in Eq. (3) reflects that a country's export basket contains multiple commodities, and operating with high market shares in product markets where industry structure closer to a monopoly indicates higher market power, which we expect to be associated with a lower commodity price elasticity.

Inflation targeting (IT). Since the 1990s, a number of central banks in both developed and developing economies have adopted inflation targeting (IT) as their monetary policy regime. Many commodity exporters have followed this trend as well. According to Roger (2009), the following fifteen commodity-exporting countries in our sample have adopted inflation targeting over the sample period (with the year of adoption in the parentheses): Australia (1993), Brazil (1999), Canada (1991), Chile (1999), Colombia (1999), Ghana (2007), Guatemala (2005), Iceland (2001), Indonesia (2005), New Zealand (1990), Peru (2002), Philippines (2002), South Africa (2000),

[10] To calculate the global supply of commodities, we use export information of 12 oil-dependent as well as 51 non-oil commodity producing countries in the world. Results are similar when commodity export information of 201 countries is used.

Thailand (2000), and Turkey (2006). In our regression specifications, dummy variable IT_t is set to unity for all the years a country operated under an inflation-targeting regime with a floating exchange rate.

Table 2 provides the summary statistics for the key variables used in our regression analyses. Mean values, in particular, will be useful in interpreting our estimation results in section 4.

INSERT TABLE 2 HERE

3.2. The baseline regression specification

In order to test our hypotheses, we consider the panel regression model commonly used in the commodity currency literature, augmented by interaction terms. That is, for country *i* and time *t*, we estimate the following:

$$REER_{it} = c_i + \beta_1 RCP_{it} + RCP_{it}\boldsymbol{X}_{it}\boldsymbol{\beta}_2 + \boldsymbol{X}_{it}\boldsymbol{\beta}_3 + \varepsilon_{it} \tag{4}$$

where both $REER$ and RCP are in logs, c_i is the country fixed effect, $\boldsymbol{X}_{it}$ is a vector of variables of interest including market power, MP, and inflation targeting, IT, which can influence the size and sign of the commodity price elasticity, and ε_{it} is a normal i.i.d. error term.

The main parameter of interest for us is on the coefficient vector $\boldsymbol{\beta_2}$, which measures how the commodity price elasticity changes in response to the structural/policy factors in $\boldsymbol{X}$. Specifically, we are interested in the following:

$$\frac{\partial[REER_{it}|RCP_{it},\boldsymbol{X}_{it}]}{\partial RCP_{it}} = \beta_1 + \boldsymbol{X}_{it}\boldsymbol{\beta}_2 \tag{5}$$

Eq. (5) shows that the magnitude of commodity price elasticity of real exchange rate depends on $\boldsymbol{X}_{it}$. A significant and positive coefficient $\boldsymbol{\beta}_2$ indicates that the higher the conditional factors in $\boldsymbol{X}_{it}$, the stronger the effect of a RCP change on $REER$.

We note that in Eq. (5), β_1 represents the commodity price elasticity when the conditioning vector $\boldsymbol{X}_{it} = 0$. Since this is not always true in the data (e.g. for the market power index we constructed), to facilitate interpretations, we apply a mean-centering approach to the continuous variables RCP_{it} and MP_{it} in the interaction term as below:

$$REER_{it} = c_i + \beta_1 RCP_{it} + \beta_2 (RCP_{it} - \overline{RCP}_i)(MP_{it} - \overline{MP}_i) + \beta_3 MP_{it} + \varepsilon_{it} \qquad (6)$$

where $\overline{RCP}_i = (1/T)\sum_{t=1}^{T} RCP_{it}$ and $\overline{MP}_i = (1/T)\sum_{t=1}^{T} MP_{it}$ for each country i. Subtracting country-specific means can also prevent the interaction terms from spuriously capturing country-varying slopes of the main term (RCP in our case) as demonstrated in Balli and Sørensen (2013). With mean-centering, β_1 now reflects the conditional effect of RCP on $REER$ when MP_{it} is at its average value for each country.

3.3. Time-series and cross-sectional properties of the data

As an illustrative step, we show in Fig. 2 the time series plots of $REER$ and RCP using a small set of countries from our sample.[11] Visual inspection of the figures shows that $REER$ and RCP are likely unit-root processes, and despite wild fluctuations individually, the two series appear to co-move over a longer period of time. There may also be structural shifts in the relationship between $REER$ and RCP. Columns (8) and (9) in Table 10 in Online Appendix A show structural

[11] The full set of countries and their major commodity export products are reported in Table 14 in Online Appendix B.

break test results based on the Gregory-Hansen (1996) cointegration test, and the selected shift dates are largely consistent with major economic events within the country. For example, we detect a structural shift in Thailand with the break date of 1997, which corresponds to the collapse of the Thai baht during the Asian financial crisis, de-coupling it from its CTOT (see the dashed vertical line in Fig. 2).

INSERT FIGURE 2 HERE

In addition to non-stationarity and cointegration, cross-sectional dependence is likely to be present in our data because common global shocks, such as world business cycles and cross-country spillovers across trading partners, may affect real exchange rates of multiple countries at the same time. By construction, $REER$'s incorporate information across trading partners. As emphasized in Driscoll and Kraay (1998), spatial correlations in panel regressions can induce a bias in standard error estimates under conventional co-variance matrix estimation techniques such as the OLS, White, and Rogers, and make statistical inferences invalid.[12]

Applying standard tests in the literature, Table 11 in the Online Appendix A summarizes test results for the presence of cross-sectional dependence, unit-root in $REER$ and RCP series, and their co-integration at conventional significance levels. We report Driscoll and Kraay (1998) standard errors to correct for spatial correlation, autocorrelation, and heteroskedasticity in our panel regressions in section 4.

[12] Ignoring cross-sectional dependence typically leads to the downward biased estimate of standard errors and the over-rejection of the null hypotheses.

3.4. Long-run vs. short-run estimation strategies

3.4.1. Long-run estimation: dynamic OLS

Due to the presence of cointegration, we use DOLS (Dynamic Ordinary Least Squares) to estimate the cointegrating vectors.[13] To improve efficiency in estimating the long-run cointegrating relationship, the DOLS estimator requires the inclusion of both the level and the leads and lags of the first difference of each regressor. Our long-run empirical specification is thus as follows:[14]

$$
\begin{aligned}
REER_{it} = c_i + \beta_1 RCP_{it} + RCP_{it}\boldsymbol{X}_{it}\boldsymbol{\beta}_2 + \boldsymbol{X}_{it}\boldsymbol{\beta}_3 \\
+ \textstyle\sum_{j=-1}^{1}\left[\gamma_{1j}\Delta RCP_{i,t+j} + \Delta\left(RCP_{i,t+j}\boldsymbol{X}_{i,t+j}\right)\boldsymbol{\gamma}_{2j} + \Delta\boldsymbol{X}_{i,t+j}\boldsymbol{\gamma}_{3j}\right] + \varphi_i t + u_{it} \quad (7)
\end{aligned}
$$

where Δ is the first-difference operator, and vector $\boldsymbol{X}_{it}$ includes MP and IT, as well as other controls when appropriate (see Section 4.3.1), c_i and $\varphi_i t$ capture country-specific fixed effect and linear trends, and u_{it} is an i.i.d. disturbance term.

3.4.2. Short-run estimation: error correction model

To examine the determinants of short-term commodity price elasticity, we represent the cointegrated system above in its error correction model (ECM) form. We look at how changes in

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[13] An alternative methodology used in the panel data analysis with non-stationary sample is FMOLS (Fully Modified Ordinary Least Squares). Kao and Chiang (2000) compare the performance of panel FMOLS and DOLS, and report that DOLS is superior in removing a finite sample bias associated with endogeneity as well as serial correlation. Note also that FMOLS requires a balanced panel, and our estimation has to rely on a substantially reduced sample size. For these reasons, we adopt DOLS as our main long-run estimation strategy.

[14] Note that interaction terms are constructed after demeaning each of the continuous variables (e.g. RCP and MP).

$REER$ respond to changes in RCP, and again our goal is to see whether any of the structural or policy measures influence this short-run relationship. The model takes the following form:

$$\Delta REER_{it} = c_i + \sum_{j=0}^{1}\left(\delta_j \Delta RCP_{i,t-j}\right) + \Delta RCP_{it}\boldsymbol{Z}_{it}\boldsymbol{\theta}_1 + \boldsymbol{Z}_{it}\boldsymbol{\theta}_2 + \lambda EC_{i,t-1} + \rho\Delta REER_{i,t-1} + \epsilon_{it} \tag{8}$$

where vector $\boldsymbol{Z}_{it}$ includes ΔMP_{it} and IT_{it}, as well as other control variables (see Section 4.3.1). The error correction term, $EC_{i,t-1}$, is the lagged residual from the first-stage estimation of cointegrating regression, $EC_{it} = REER_{it} - c_i - \varphi_i t - \hat{\beta}_1 RCP_{it}$.[15]

4. Empirical results

4.1. Impact of market power on the long-run elasticity

Table 3 reports estimates for the long-run DOLS specification in Eq. (7). From the first row, we note that there is a strong and robust connection between the $REER$ and RCP. The panel regressions show a roughly one-for-one or slightly stronger co-movement in the long-run between exchange rates and commodity prices. The range of estimates is consistent with results reported in prior literature.

INSERT TABLE 3 HERE

[15] We also test validity of our empirical specifications given in Eqs. (7) and (8) by employing least angle regressions (LARS) of Efron et al. (2004). This model selection method strongly supports our empirical model specifications for both long-run and short-run (see Online Appendix C for details). Moreover, the panel Granger non-causality test of Dumitrescu and Hurlin (2012), results available in Table 12 in Online Appendix A, supports a unidirectional causality from ΔRCP to $\Delta REER$ and justifies our single-equation error correction model.

We next turn to the interaction effects between country-specific commodity prices and our variables of interest $\boldsymbol{X}$, to see how the structural and policy variables induce differential commodity price elasticities. The negative coefficient estimate for $RCP_t \times MP_t$ in column (2) shows that as a country's market power increases, $REER$ reacts *less* to a given RCP change; this is consistent with the hypothesis we put forth in Section 2.1. Specifically, the long-run commodity price elasticity is estimated to be 1.27 for countries with a period-average level of MP, indicating that a 1% increase in RCP is associated with a 1.27% increase in $REER$. If a country has an MP index that is one unit above its period-mean, the -1.26 coefficient estimate indicates that the country is expected to have 1.26% lower elasticity.

In regression specifications involving the IT dummy variable, a choice needs to be made regarding the omitted reference category in order to focus on the performance of IT regime only; we set it to be the group of countries with no explicit commitment to either domestic inflation or exchange rate controls, i.e. the non-IT floaters. We then take two approaches. The first approach, reported in columns (3) and (4), eliminates observations and uses sub-samples: countries that have ever adopted a hard peg in any part of the sample period, including some that became IT countries, are excluded.[16] The second approach uses the full sample but controls for exchange rage pegs, as reported in columns (5) and (6).[17] The results are very similar.

[16] These include IT countries such as Guatemala, Philippines, and Thailand. Moreover, observations in years 1980 and 1981 were eliminated because Chile, Madagascar, and Pakistan operated under a fixed exchange rate regime during those periods. This explains why columns (3) and (4) of Table 3 have 23 countries, covering 1982-2010.

[17] Note that we classify a country into the category of peggers if it takes a *de facto* peg or pre announced band with margins of no larger than +/-2% based on the fine classification of Ilzetzki et al. (2017).

Estimates in column (3) show that the adoption of inflation targets has no effect on the long-run commodity price elasticity. The IT coefficient is not significant either. This result is consistent with the standard neoclassical assumption of long-run monetary neutrality.

Including the market power interaction term in column (4) does not alter the statistical insignificance of the IT interaction term, though the negative market power effect on the real exchange rate reaction remains statistically significant. In columns (5) and (6), we see that the full-sample results with a hard currency peg (PEG) control are consistent with findings from the sub-sample specifications. In addition, we see some evidence for a more muted long-run commodity price elasticity in countries with fixed exchange rates, compared to the reference-group countries that have flexible exchange rates and no inflation targets.

In addition to individual coefficient estimates, the table also reports F-statistics for the null hypothesis that RCP has no effect on $REER$ in the specifications that include interaction effects. As seen in Eq. (5), this null hypothesis requires a joint significance test for β_1 and $\boldsymbol{\beta}_2$. Table 3 shows that the *p*-values for the F-statistics are all consistently below 1%, indicating that that RCP and the interaction terms are jointly significant and informative in explaining long-run $REER$ behavior.

4.2. Impact of inflation targeting on the short-run elasticity

We now investigate the workings of commodity currencies in the short-run, following the error correction specification discussed in the previous section. Table 4 displays the parameter estimates from the second stage of the ECM specification or Eq. (8).

INSERT TABLE 4 HERE

Across the first row in Table 4, we see a clear relationship between changes in $REER$ and changes in RCP. The range of elasticity estimates for contemporaneous RCP movements suggest a more gradual adjustment of $REER$, consistent with the possible presence of short-run rigidities. We also note that all estimates of the error-correction term (EC) have the expected negative signs at the l% level of statistical significance, indicating the presence of a long-run cointegrating relationship. Interestingly, the estimated error-correction coefficient of -0.21, reported in column (1), for annual adjustment towards the long-run equilibrium has an implied half-life of close to three years, similar to the three to five year consensus (Rogoff, 1996) for the typical half-lives of real exchange rate to shocks.

We also find that lagged commodity price movements, ΔRCP_{t-1}, can lead to dynamic responses in real exchange rates in our panel data, in contrast to findings in some earlier studies. For example, Chen and Rogoff (2012) use local-to-unity analysis without considering the error-correction mechanism in their predictive regressions. Using data from a much smaller set of countries, they find no consistent dynamic relation between commodity price movements and subsequent real exchange rate behavior. The ECM results based on our larger set of panel data suggest that commodity prices may indeed have predictive power for future real exchange rate movements.

Turning to our parameter estimates of interest, the negative and statistically significant coefficient of the MP_t interaction term in column (2) suggests that even in the short-run, a higher market power may reduce the real exchange rate response to the commodity price shocks. This result, however, was not robust in the other specifications.

Unlike in the long-run results, we see that IT tends to amplify the impact of RCP shocks on $REER$ in the short-run. This suggests that domestic policies designed to stabilize medium-term inflation may come at the cost of creating excess short-term exchange rate volatility in the face of CTOT shocks, whose transmission into the real exchange rate would stem mostly from the nominal exchange rate when prices remain sticky. In fact, the results from columns (3) and (4) indicate that the short-run elasticity seems to be about 5 times larger in IT countries than non-IT countries.

Returning to the full sample with a hard currency peg controlled, the messages emerging from columns (5) and (6) are essentially the same, except that we no longer see any significant short-term impact from the market power interaction term. The fixed exchange rate regime interaction term shows a significantly negative sign as expected. Our result regarding the effect of the exchange rate regime generally supports Broda (2004)'s finding that the real exchange rate appreciation in response to a rise in terms-of-trade is small and slow in pegs while it is large and immediate in floats.

One final remark is that the overall goodness-of-fit is generally much lower in the short-run than the long-run regressions, as also suggested by LARS procedures in Online Appendix C.

4.3. Robustness checks

4.3.1. Other potential elasticity determinants

Admittedly, this is not the first paper to explore potential determinants of the commodity price elasticity of the real exchange rate. As mentioned in the introduction, Aizenman et al. (2012) emphasized a cushioning role of international reserves on the short-run elasticity. Bodart et al. (2012, 2015) pointed to the differential impact of export diversification, exchange rate regime, and trade and financial openness on the long-run elasticity. In this subsection, we test how sensitive our main results are when these other elasticity determinants are also controlled in the baseline regression specifications. Before presenting the relevant results, we motivate their inclusions based on standard theory and discuss their expected signs below:

Trade openness (TO). The sum of exports and imports relative to GDP is used to measure how much the economy relies on tradable goods (source: WB WDI). The small open economy model in the commodity currency literature predicts that price changes in commodity exports are transmitted into the real exchange rate primarily through adjustments in the non-traded goods prices. As such, we expect a country with a larger tradable sector (or more open to trade) to have a more muted real exchange rate response to CTOT movements.

Financial openness (FO). We use the Chinn-Ito (2006) index as a *de jure* measure of capital account openness; higher values indicate more open capital markets. Our hypothesis is that the commodity price elasticity may be smaller in the presence of cross-border capital control. Theoretically, this is because inelastic supply of capital in response to the commodity boom would limit the rise in wages and prices in the non-traded goods sector. This suggests that a stronger real exchange rate response is expected in countries that are more open to cross-border capital transactions.

Commodity export dependency (CEX). Commodity export dependency reflects how diversified exports are in a commodity-producing country, and is proxied by the ratio of the country's total commodity exports to its total goods exports (source: UN COMTRADE). Accordingly, a higher value of CEX implies heavier reliance on commodity exports, and it is likely to strengthen the commodity currency phenomenon.

International reserves (RES). A ratio of total reserves (excluding gold) to nominal GDP is included to capture the standard view that a sizeable stock of international reserves enables a central bank to counteract exchange rate fluctuations more easily through foreign exchange interventions (source: IMF IFS and WB WDI). As such, one would expect the commodity price elasticity estimate to be smaller in countries with large international reserves and engage in active foreign exchange management.

Hard currency peg (PEG). In order to evaluate the effect of a country's choice to adopt a hard currency peg regime, we follow Ilzetzki et al. (2017) and adopt their detailed classification scheme, where a larger fine classification code corresponds to a more flexible nominal exchange rate regime. We create a dummy variable PEG_t which takes the value of unity at time t if the fine classification code for the country is less than five, and zero otherwise. The commodity price elasticity is expected to be smaller under fixed exchange regimes when we assume nominal rigidity is present. However, one cannot exclude the possibility that a hard peg can amplify the transmission of commodity price shocks when prices become flexible in the long-run.

Table 5 summarizes the results when these additional determinants and their interaction terms are introduced to our baseline models. Since the inclusion of all variables, with leads and lags as well as fixed effects, involves a large number of regressors and possible collinearity among them, the power of the statistical inference may lessen for some of the estimates, given our sample size. Nevertheless, the main message remains the same: we see statistically significant coefficients for the MP interaction term in the long-run and for the IT interaction term in the short-run, both with the expected signs.

Concerning the globalization variables, from column (1) in Table 5, we see that trade and financial openness both have a significant impact on the long-run elasticity, with signs consistent with our theoretical predictions. The long-run elasticity is smaller for countries with more open international trade or stricter cross-border capital controls.[18] In the short-run, shown in column (2),

[18] Interestingly, Bodart et al. (2015) find the opposite empirical result for the effect of financial openness on the long-

these openness measures still affect the elasticity. However, the signs of the openness interaction terms are reversed. Last but not least, large international reserves and currency pegs appear to decrease the short-run elasticity, while neither of them has a significant impact on the long-run elasticity.

INSERT TABLE 5 HERE

One policy implication from our ECM results in Table 5 column (2) is as follows. Emerging market economies that have adopted IT around 2000 (e.g. Brazil, Chile, Colombia, Indonesia, Peru, Philippines, South Africa, Thailand, and Turkey) typically hold larger stockpiles of international reserves as a share of GDP, as illustrated in Online Appendix D Fig. 3. Accumulating massive reserves, as pointed out in prior work, may thus reflect these countries' desire to carve out a particular niche in the policy trilemma space: ultimately, they want to achieve an intermediate level of monetary independence, some degree of exchange rate stability in the presence of volatile global commodity prices, and at the same time retaining global financial integration.[19]

4.3.2. Structural shifts in cointegration

Next, we consider the potential structural shifts in the long-run cointegration, as indicated in Fig. 2 in the case of Thailand. Abrupt changes in $REER$ would blur the cointegrating

run elasticity, which is also in contrast to their small open economy model prediction.

[19] See Aizenman et al. (2010, 2013) among others for the emerging market countries' convergence towards a middle ground in the trilemma space by relaxing trilemma constraints through active foreign exchange intervention. In the context of emerging market inflation targeters, Ghosh et al. (2016) also underscore a supplementary role of the foreign exchange intervention that may further enhance the credibility of the inflation targeting framework.

relationship between $REER$ and RCP and need to be properly controlled. To this end, we adopt a cointegration test proposed by Gregory-Hansen (1996) that allows for regime shifts at an unknown point in time. Formally, the Gregory-Hansen test is based on the model below:

$$REER_t = \alpha_0 + \alpha_1 \vartheta_{t\pi} + \alpha_2 RCP_t + \varepsilon_t \tag{9}$$

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

$$\vartheta_{t\pi} = \begin{cases} 0 & \text{if } t \leq [N\pi] \\ 1 & \text{if } t > [N\pi] \end{cases} \tag{10}$$

where N is the sample size, $\pi \in (0,1)$ determines the timing of the level shift, and [ ] rounds the input to an integer. Following Gregory and Hansen (1996), we choose a trim (π) of 0.15, which specifies the fraction of the data range that is skipped at either ends when testing for possible break points. The test is applied to each country to detect possible shift dates, while the lag length is chosen based on the Akaike information criterion (AIC). The null hypothesis tests for no cointegration against the alternative of cointegration with a single shift at an unknown point in time.

Columns (8) and (9) of Table 10 in Online Appendix A report the Gregory-Hansen $Z(t)$ statistics and shift dates selected by the test at the 5% level of statistical significance. The selected shift dates are largely consistent with the countries' major macroeconomic events, such as hyperinflation, exchange rate crisis, and nominal exchange rate reforms.

We then account for the regime shifts in the cointegrating relationship by including country-level structural shift dummy variables in the DOLS(1,1) and first-stage ECM specifications. In Table 6, we find that our earlier results remain robust: there is a negative MP interaction effect in the long-run and a positive IT interaction effect in the short-run.

INSERT TABLE 6 HERE

4.3.3. Major international financial crises

Exposure to severe financial turbulence that brings about large swings in the global markets can significantly affect the currency values of the afflicted countries. Therefore, major financial crisis events have the potential to obscure the relation between $REER$ and RCP. Since the Gregory-Hansen (1996) test above is designed to identify a single level shift, it may not appropriately cover the length of time during which certain countries in our sample were affected by global turmoil. For this reason, our next robustness test includes dummy variables that control for the following major international financial crises: 1982-83 Latin American debt crisis (Argentina and Brazil), 1997-98 Asian crisis (Indonesia, Malaysia, Philippines, and Thailand), 2001-02 Argentine crisis, and 2008-09 Global Financial Crisis.

We see from the results in Table 7 that the main finding remains stable and significant when the effects of major financial crises are accounted for.

INSERT TABLE 7 HERE

4.3.4. Potential endogeneity issues

One of the main objectives of this paper is to examine whether the strength of the commodity currency phenomenon is altered when the small open economy assumption is relaxed. Deviating from this standard price-taker assumption implies that it is possible for commodity prices to be endogenous in the real exchange rate regressions due to reverse causality.

Our long-run and short-run estimates based on cointegration methods are unlikely to be affected by a simultaneity bias. The DOLS estimator uses a parametric correction for endogeneity by including leads and lags of the differenced RCP variable. In regards to the ECM estimates, the panel Granger non-causality test of Dumitrescu and Hurlin (2012) ascertains uni-directional causation from commodity prices to the real exchange rate.[20]

As an additional confirmation, we test the robustness of our main results to treating RCP as endogenous. For consistent estimations of our long-run model, we employ a two-step efficient GMM estimator using the two- and three-year lagged values of RCP as joint instruments for contemporaneous RCP, including its value in the interaction terms. For the ECM estimation, we apply the two-step system GMM estimator of Arellano and Bover (1995) or Blundell and Bond (1998), with Windmeijer (2005) small sample robust correction. The system GMM estimator uses appropriately lagged values of the levels and differences of the suspect endogenous variables as instruments to address any potential endogeneity arising from RCP and interaction terms as well as from lagged changes in $REER$.[21]

Table 8 presents the regression results using the instrumental-variables (IV) approach. For the long-run static model estimation in columns (1)-(3), we first confirm the overall validity of the instruments. First, Hansen's J statistics do not reject the null hypothesis that the over-identifying restrictions are valid (i.e. the IV's are orthogonal to the disturbance term). Additionally, weak instruments do not appear to be a concern, as signaled by the Kleibergen-Paap (2006) rank Wald

[20] Moreover, the endogeneity test that relies on a *C* statistic, defined as the difference of two Sargan/Hansen statistics between regressions with or without treating RCP as endogenous within the two-step efficient generalized method of moments (GMM) procedure, fails to reject its null hypothesis at the 95% level. This evidence suggests that RCP may be treated as exogenous (results available in Table 16 in Online Appendix E).

[21] Efficiency gain of the system GMM comes from the introduction of changes in the instrument variables that are assumed to be uncorrelated with the fixed effects. Following a traditional Hausman and Taylor (1981) approach, we could use deviations of time-varying variables from their within-group means as instruments in order to address the endogeneity with respect to the unobserved unit-specific effects. Our main result is robust to this alternative specification whose estimates are available in Table 17 in Online Appendix E.

statistic, which well exceeds the rule-of-thumb threshold of ten suggested by Staiger and Stock (1997). For the short-run dynamic model estimation in columns (4)-(6), the non-rejection of the Hansen over-identification and second-order autocorrelation tests gives assurance that the set of instruments is appropriate.

INSERT TABLE 8 HERE

The estimated coefficients for RCP, ΔRCP, and their interaction terms under static and dynamic GMM estimations are shown in Table 8. They all stay within a reasonable range from the main results reported earlier, as well as showing a consistent pattern of statistical significance. Having more market power in the world commodity market tends to dampen the transmission of commodity price changes into the real exchange rate in the long-run, and inflation-targeting regimes tend to amplify it in the short-run.[22] These findings provide reassurance that the presence of potential commodity price endogeneity is not a major concern behind our main cointegrating regression results.[23]

4.3.5. Omission of outliers

Looking closely at Fig. 1, one may notice the presence of outliers in our sample. The estimates of long-run elasticity in Ghana (7.15), Mozambique (10.93), and Paraguay (6.20) seem exceptionally large compared to the median value of 2.32 (standard deviation is 2.98). One may

[22] Unreported one-step difference GMM estimator produces qualitatively similar short-run results.

[23] Another potential source of bias may arise from the presence of dynamic heterogeneous panels. We address it using a pooled mean group (PMG) estimator of Pesaran, Shin, and Smith (1999), which allows a cross-country heterogeneity in the short-run coefficients including the speed of adjustment while imposing long-run slope homogeneity. Overall, the PMG estimates seem weaker than the main ECM results, but they become qualitatively similar to the baseline results once we control for cross-sectional dependence in residuals (see Table 18 in Online Appendix E).

worry that these outliers may be driving our earlier findings, so we exclude them as another robustness check. Table 9 demonstrates that our main conclusions stay unchanged when these outliers are excluded from the sample.

INSERT TABLE 9 HERE

5. Conclusion

This paper studies the real exchange rate behavior for a large set of natural resource-abundant developing economies. Due to these countries' heavy commodity-export dependency, their real exchange rates react strongly to changes in the world price of their commodity products. According to our empirical evidence based on various panel data estimations, the real exchange rates of these countries appreciate when the global commodity prices rise, both in the short-run and in the long-run.

Persistent real appreciations and sharp real deprecations both have potential detrimental impacts on the local economy, such as the loss of international competitiveness for the domestic manufacturing industries, or the heavier debt burden of local firms with liabilities denominated in foreign currencies. For these reasons, exchange rate stabilization is a major policy concern in commodity-producing developing countries.

This paper contributes to understanding of cross-country variations in real exchange rate responses to global commodity price shocks. Focusing on the role of market structure and monetary policy channels, amongst others, the paper finds that in the long-run, larger market power in the world commodity market can reduce the exchange rate response, while in the short-run, an inflation targeting regime can amplify it. The result also points to, in line with previous findings, the importance of globalization strategies in both goods and financial markets, and of foreign exchange intervention in counteracting commodity currency fluctuations.

Appendix

A theoretical framework for market power and commodity currencies

In order to illustrate the theoretical linkage between a country's market power in the world commodity market and the commodity price elasticity of its real exchange rate, we outline a structural model that relaxes the pure price-taker assumption commonly assumed for small open economies. The model below is a simplified version of the two-country supply-demand model presented in Swift (2004) and Clements and Fry (2008), where the home country is assumed to produce only commodities and export them all to the foreign country.

The commodity supply (Y_c^s) and demand (Y_c^d) at home and abroad are assumed to depend on the relative prices as follows:

$$Y_c^s = Y_c^s\left(\frac{p_c}{P}\right); \qquad Y_c^d = Y_c^d\left(\frac{p_c^*}{P^*}\right) \tag{A1}$$

where p_c is a commodity price in a domestic currency, P is consumer price index in the home country, and an asterisk superscript denotes a foreign variable. The law of one price is assumed for the tradable commodities only, $Ep_c = p_c^*$, where E is the nominal exchange rate, defined as the price of domestic currency in terms of foreign currency.

Purchasing power parity for the overall consumption baskets is not imposed, so the rate of change of the real exchange rate (Q, the relative price of domestic consumption basket in terms of foreign consumption basket) can be expressed as below:

$$\hat{Q} = \hat{E} + \hat{P} - \hat{P}^* \tag{A2}$$

where a hat above a variable indicates a logarithmic derivative.

As in Clements and Fry (2008), log-differentiating Eq. (A1) and imposing the market clearing condition for the commodities, we can express real exchange rate change in Eq. (A2) as:

$$\hat{Q} = \left(\frac{\varepsilon^s - \varepsilon^d}{\varepsilon^s}\right)\left(\frac{\widehat{p_c^*}}{P^*}\right) = \eta\left(\frac{\widehat{p_c^*}}{P^*}\right) \tag{A3}$$

where ε^s (≥ 0) is the price elasticity of commodity supply and ε^d (≤ 0) is the price elasticity of commodity demand. We see that an increase in the real world price of commodities leads to a real appreciation of the home currency, with the magnitude determined by the parameter $\eta \geq 1$, which increases with the supply elasticity and decreases with the demand elasticity (a negative number). In other words, more elastic demand curves correspond to larger real exchange rate responses to a given commodity price change.

Our inference is based on the assumption that countries with higher global market shares are likely to face less elastic demand curves, and the higher ε^d (smaller in absolute value) would then imply a smaller η and a weaker commodity-currency response. Market sizes may also affect the supply elasticity, possibly through economies-of-scale in production, fixed cost to export-market entry, and so forth. That is, our market power (MP) variable can proxy η through both ε^d *and* ε^s, even though our main story emphasizes the more natural demand channel (so strictly speaking, we are testing for a *net* market power or demand effect.)

We note that the discussion above does not require these commodity-producing countries to be global monopolists or have substantial market power like the OPEC countries, and merely that they do not face absolutely flat demand curves or are pure price-takers. Our null hypothesis is that relatively high market share, everything else equal, should weaken the exchange rate response.

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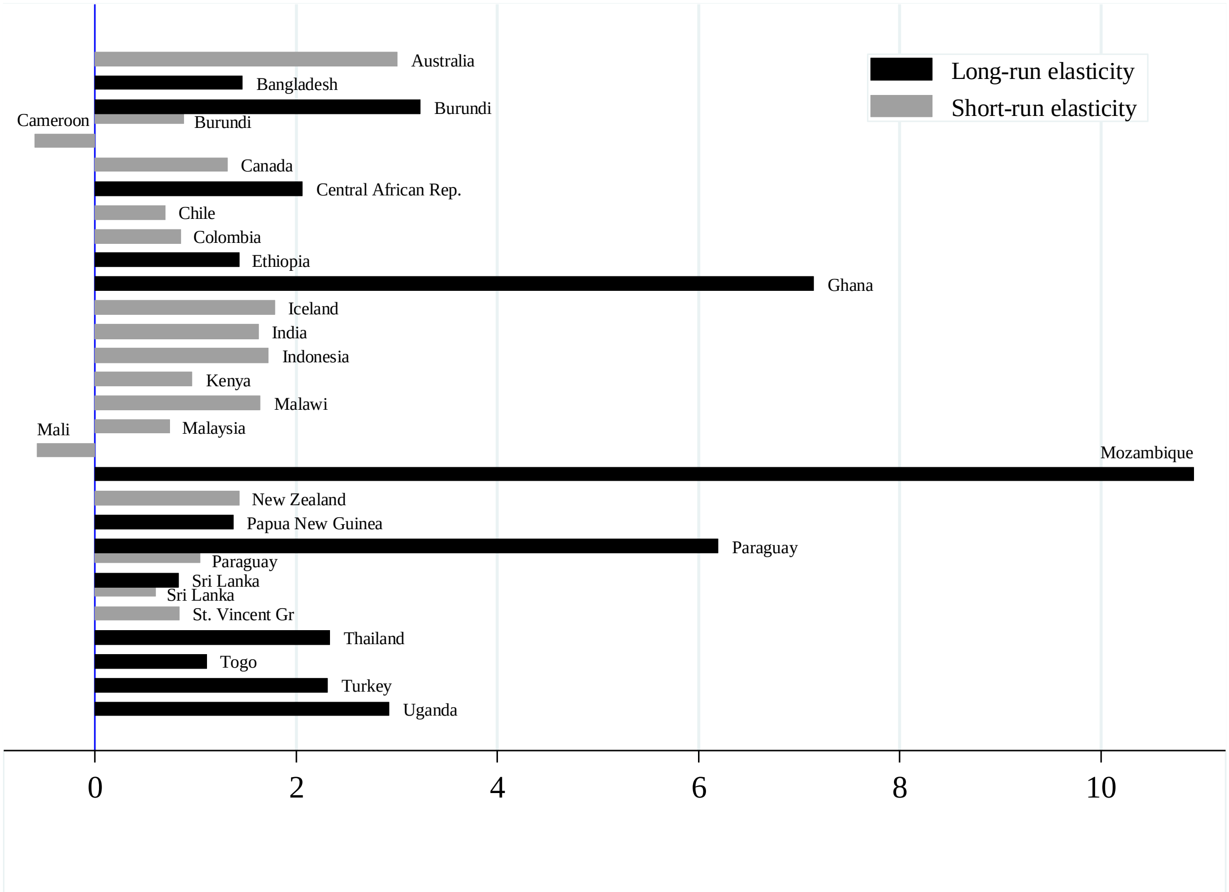

Fig. 1. Distribution of the commodity price elasticity of real exchange rates across countries. Notes: We estimate the long-run elasticity, β_1, by the dynamic OLS ($REER_t = \beta_0 + \beta_1 RCP_t + \sum_{j=-1}^{1} \gamma_j \Delta RCP_{t+j} + u_t$), and the short-run elasticity, δ_1, by the first differencing ($\Delta REER_t = \delta_0 + \delta_1 \Delta RCP_t + \epsilon_t$) based on the unit-root and cointegration test results given in Table 10 in Online Appendix A. Real commodity prices (RCP) and real effective exchange rates ($REER$) are measured in logarithm.

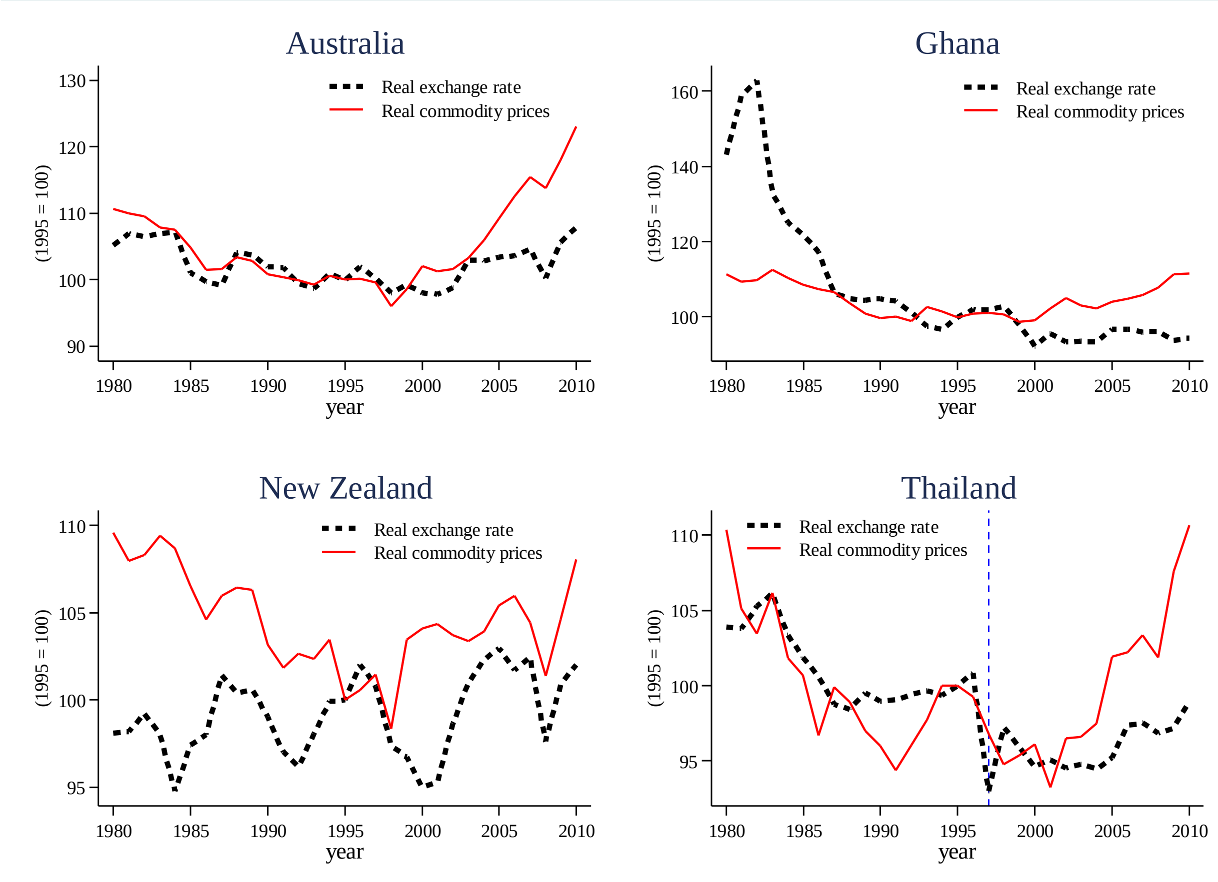

Fig. 2. Time series plots of the real exchange rate and real commodity prices. Notes: The thin and dashed vertical line for Thailand indicates a structural shift date.

Table 1
Commodities with 25% or more world export share.

Commodity	Exporting country	Average world export share over the period			
		1980-1990	1991-2000	2001-2010	1980-2010
		(1)	(2)	(3)	(4)
Agricultural food					
Cocoa	Cote d'Ivoire	47.7	49.6	39.3	44.1
Coconut oil	Philippines	66.8	58.2	49.4	58.1
Palm kernel oil	Indonesia	13.6	40.1	54.2	39.4
Palm kernel oil	Malaysia	76.7	49.3	33.5	53.9
Palm oil	Malaysia	73.9	65.9	47.4	62.8
Rice	Thailand	30.0	27.9	26.7	28.2
Soybean meal	Argentina	18.2	23.2	35	28.1
Soybean meal	Brazil	38.8	29.1	24.2	29.4
Soybean oil	Argentina	14.0	29.7	41.9	28.0
Sunflower oil	Argentina	45.7	37.7	20.7	35.0
Ag. raw materials					
Hard Logs	Malaysia	61.5	35.6	16.7	38.7
Softwood sawn	Canada	45.3	45.7	30	40.5
Wood pulp	Canada	33.5	29.4	20.3	27.9
Wool, coarse	Australia	71.1	75.6	73.9	73.4
Wool, fine	Australia	32.4	41.6	31.7	35.2
Wool, fine	New Zealand	39.0	32.4	35.6	35.8
Fertilizers					
Phosphate rock	Morocco	43.9	39.8	41.7	41.9
Metals					
Copper	Chile	32.6	31.3	36.8	33.6
Iron	Australia	25.7	24.1	29.9	26.6
Iron	Brazil	33.2	32.2	28.7	31.2
Nickel	Canada	31.1	24.8	19.0	25.2
Uranium	Niger	51.4	33.0	26.0	31.2
Uranium	Australia	60.5	59.7	48.4	56.3
Energy					
Coal	Australia	25.2	28.7	30.6	28.1

Notes: Commodities whose average world export shares larger than 25% over the period 1980-2010 are documented only. Columns (1)-(3) display decade-by-decade statistics and column (4) reports average shares during the whole sample period. Source: All figures are calculated by authors using the UN COMTRADE data.

Table 2
Summary statistics for key variables.

Variable	Observations	Mean	Std. Dev.	Min	Max
$REER$	1,581	4.747	0.371	3.362	7.841
$\Delta REER$	1,530	-0.014	0.161	-1.438	1.223
RCP	1,581	1.293	0.446	0.357	3.917
ΔRCP	1,530	0.001	0.046	-0.323	0.435
MP	1,271	0.497	0.922	0	6.600
ΔMP	1,164	-0.007	0.176	-1.141	3.512
IT	1,581	0.105	0.307	0	1

Notes: $REER$ = real effective exchange rate; RCP = real commodity price; MP = market power; IT = inflation targeting.

Table 3
Long-run commodity price elasticity and interaction effects.

Estimation method:	Dynamic OLS(1,1)					
Dependent variable:	$REER_t$					
	(1)	(2)	(3)	(4)	(5)	(6)
RCP_t	**1.30*****	**1.27*****	**1.13****	**1.77*****	**1.17*****	**1.60*****
	(0.20)	(0.26)	(0.27)	(0.46)	(0.26)	(0.45)
$RCP_t \times MP_t$		**-1.26*****		**-1.16****		**-1.61*****
		(0.38)		(0.44)		(0.39)
$RCP_t \times IT_t$			**0.46**	**-0.85**	**0.56**	**-0.45**
			(0.35)	(0.69)	(0.38)	(0.61)
$RCP_t \times PEG_t$					-0.08	-1.36**
					(0.31)	(0.54)
MP_t		0.06**		-0.004		0.08***
		(0.03)		(0.02)		(0.02)
IT_t			0.11	0.05	0.13**	0.07
			(0.09)	(0.06)	(0.06)	(0.05)
PEG_t					0.17***	0.10***
					(0.03)	(0.03)
R^2	0.47	0.46	0.44	0.47	0.50	0.48
F-statistics		15.14***	14.04***	9.20***	20.19***	15.58***
Sample period	1980-2010	1980-2010	1982-2010	1982-2010	1980-2010	1980-2010
# countries	51	51	23	23	51	51
Observations	1,428	1,012	644	482	1,428	1,012

Notes: $REER$ = real effective exchange rate; RCP = real commodity price; MP = market power; IT = inflation targeting; PEG = hard currency peg. With exclusion of countries with hard peg regimes, the subsample in columns (3) and (4) includes only 23 countries such as Australia, Brazil, Canada, Chile, Colombia, Ethiopia, Ghana, Iceland, Indonesia, Madagascar, Mauritania, Mauritius, New Zealand, Pakistan, Papua New Guinea, Peru, South Africa, Sri Lanka, Sudan, Tanzania, Turkey, Uruguay, and Zambia. DOLS(1,1) procedure includes contemporaneous, 1 lead, and 1 lag of changes of RCP and interaction terms but they are suppressed to save a space. The regressions also control for country fixed effects and time trend. Driscoll-Kraay standard errors are reported in parentheses. F-statistics for a Wald test and their significance level are reported to test the joint significance of coefficients for contemporaneous RCP and interaction terms. ***, ** indicate statistical significance at the 1% and 5% levels, respectively.

Table 4
Short-run commodity price elasticity and interaction effects.

Estimation method:	Error correction model					
Dependent variable:	$\Delta REER_t$					
	(1)	(2)	(3)	(4)	(5)	(6)
ΔRCP_t	**0.39*****	**0.47*****	**0.24***	**0.29****	**0.37*****	**0.41*****
	(0.08)	(0.11)	(0.13)	(0.13)	(0.13)	(0.13)
$\Delta RCP_t \times \Delta MP_t$		**-1.99***		**-0.41**		**-1.81**
		(1.17)		(0.79)		(1.20)
$\Delta RCP_t \times IT_t$			**1.07*****	**1.05*****	**0.84*****	**0.79*****
			(0.25)	(0.25)	(0.21)	(0.21)
$\Delta RCP_t \times PEG_t$					-0.29*	-0.24**
					(0.16)	(0.11)
ΔMP_t		0.005		-0.04**		0.002
		(0.05)		(0.02)		(0.05)
IT_t			0.01	0.02	0.03	0.03
			(0.02)	(0.03)	(0.02)	(0.02)
PEG_t					0.05***	0.04***
					(0.01)	(0.01)
EC_{t-1}	-0.21***	-0.19***	-0.26***	-0.26***	-0.21***	-0.20***
	(0.03)	(0.03)	(0.03)	(0.04)	(0.03)	(0.03)
ΔRCP_{t-1}	0.28**	0.21*	0.30*	0.18*	0.24**	0.16*
	(0.12)	(0.11)	(0.15)	(0.09)	(0.12)	(0.09)
$\Delta REER_{t-1}$	-0.02	-0.09*	0.06	-0.06	-0.02	-0.09*
	(0.04)	(0.05)	(0.05)	(0.05)	(0.03)	(0.05)
R^2	0.16	0.18	0.28	0.30	0.17	0.19
F-statistics		17.49***	36.96***	25.84***	26.35***	23.82***
Sample period	1980-2010	1980-2010	1982-2010	1982-2010	1980-2010	1980-2010
# countries	51	51	23	23	51	51
Observations	1,479	1,135	644	515	1,479	1,135

Notes: $REER$ = real effective exchange rate; RCP = real commodity price; MP = market power; IT = inflation targeting; PEG = hard currency peg; EC = error correction term. With exclusion of countries with hard peg regimes, the subsample in columns (3) and (4) includes only 23 countries such as Australia, Brazil, Canada, Chile, Colombia, Ethiopia, Ghana, Iceland, Indonesia, Madagascar, Mauritania, Mauritius, New Zealand, Pakistan, Papua New Guinea, Peru, South Africa, Sri Lanka, Sudan, Tanzania, Turkey, Uruguay, and Zambia. All specifications control for country fixed effects. Driscoll-Kraay standard errors are reported in parentheses. F-statistics for a Wald test and their significance level are reported to test the joint significance of coefficients for ΔRCP_t and interaction terms. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 5
Commodity price elasticity and interaction effects: controlling for other potential determinants.

Estimation method:	Dynamic OLS(1,1)		Error correction model
Dependent variable:	$REER_t$		$\Delta REER_t$
	(1)		(2)
RCP_t	**1.14***** (0.28)	ΔRCP_t	**0.46***** (0.15)
$RCP_t \times MP_t$	**-1.27***** (0.44)	$\Delta RCP_t \times \Delta MP_t$	**-0.66** (1.09)
$RCP_t \times IT_t$	**-0.08** (0.48)	$\Delta RCP_t \times IT_t$	**0.70***** (0.21)
$RCP_t \times TO_t$	-5.60*** (1.90)	$\Delta RCP_t \times \Delta TO_t$	1.30* (0.72)
$RCP_t \times FO_t$	1.06* (0.61)	$\Delta RCP_t \times \Delta FO_t$	-1.33** (0.61)
$RCP_t \times CEX_t$	0.58 (1.52)	$\Delta RCP_t \times \Delta CEX_t$	-0.27 (0.88)
$RCP_t \times RES_t$	4.47 (4.13)	$\Delta RCP_t \times \Delta RES_t$	-5.37* (3.02)
$RCP_t \times PEG_t$	-0.90 (0.58)	$\Delta RCP_t \times PEG_t$	-0.32** (0.13)
MP_t	-0.002 (0.03)	ΔMP_t	0.04 (0.05)
IT_t	0.03 (0.04)	IT_t	0.02 (0.02)
TO_t	-0.30*** (0.09)	ΔTO_t	-0.39*** (0.11)
FO_t	0.32*** (0.04)	ΔFO_t	0.03 (0.05)
CEX_t	0.36*** (0.13)	ΔCEX_t	0.03 (0.04)
RES_t	0.49*** (0.16)	ΔRES_t	-0.12 (0.15)
PEG_t	0.10*** (0.03)	PEG_t	0.03*** (0.01)
		EC_{t-1}	-0.17*** (0.03)
		ΔRCP_{t-1}	0.31*** (0.11)
		$\Delta REER_{t-1}$	-0.17*** (0.04)
R^2	0.58		0.22
F-statistics	73.33***		18.05***
Sample period	1980-2010		1980-2010
# countries	51		51
Observations	993		1,118

Notes: $REER$ = real effective exchange rate; RCP = real commodity price; MP = market power; IT = inflation targeting; TO = trade openness; FO = financial openness; CEX = commodity export dependency; RES = international reserves; PEG = hard currency peg; EC = error correction term. See notes to Table 3 and 4 for a description of the estimation methods.

Table 6
Commodity price elasticity and interaction effects: controlling for regime shift dates.

Estimation method:	Dynamic OLS(1,1)				Error correction model		
Dependent variable:	$REER_t$				$\Delta REER_t$		
	(1)	(2)	(3)		(4)	(5)	(6)
RCP_t	**1.09*****	**0.97*****	**1.48*****	ΔRCP_t	**0.48*****	**0.24***	**0.42*****
	(0.30)	(0.28)	(0.53)		(0.11)	(0.13)	(0.13)
$RCP_t \times MP_t$	**-1.22*****		**-1.54*****	$\Delta RCP_t \times \Delta MP_t$	**-1.97***		-1.81
	(0.37)		(0.42)		(1.16)		(1.19)
$RCP_t \times IT_t$		**0.33**	**-0.56**	$\Delta RCP_t \times IT_t$		**1.02*****	**0.75*****
		(0.39)	(0.77)			(0.25)	(0.21)
$RCP_t \times PEG_t$			-1.33**	$\Delta RCP_t \times PEG_t$			-0.25**
			(0.55)				(0.10)
MP_t	0.07**		0.08***	ΔMP_t	0.005		0.002
	(0.03)		(0.03)		(0.05)		(0.05)
IT_t		0.04	0.02	IT_t		0.001	0.02
		(0.07)	(0.04)			(0.02)	(0.02)
PEG_t			0.10***	PEG_t			0.04***
			(0.03)				(0.01)
				EC_{t-1}	-0.24***	-0.32***	-0.24***
					(0.03)	(0.03)	(0.03)
				ΔRCP_{t-1}	0.19*	0.29*	0.16*
					(0.10)	(0.16)	(0.09)
				$\Delta REER_{t-1}$	-0.07	0.07	-0.07
					(0.05)	(0.05)	(0.05)
R^2	0.57	0.57	0.58		0.21	0.31	0.22
F-statistics	9.92***	9.78***	8.48***		15.72***	32.29***	22.62***
Sample period	1980-2010	1982-2010	1980-2010		1980-2010	1982-2010	1980-2010
# countries	51	23	51		51	23	51
Observations	1,012	644	1,012		1,135	644	1,135

Notes: See notes to Table 3 and 4 for a description of the variable abbreviations and the estimation methods. In DOLS(1,1) and first-stage error correction model specifications, level shift dummies are included to control for structural shift dates identified by the Gregory-Hansen (1996) cointegration test (Bolivia, 1985; Burundi, 2002; Central African Republic, 1992; Costa Rica, 1986; Ethiopia, 1992; Madagascar, 1985; Malawi, 1992; Papua New Guinea, 1997; Peru, 1987; Senegal, 1992; Tanzania, 1986; Thailand, 1997; Togo, 1992; Uganda, 1988).

Table 7
Commodity price elasticity and interaction effects: accounting for international financial crises.

Estimation method:	Dynamic OLS(1,1)				Error correction model		
Dependent variable:	$REER_t$				$\Delta REER_t$		
	(1)	(2)	(3)		(4)	(5)	(6)
RCP_t	**1.28*****	**1.04*****	**1.59*****	ΔRCP_t	**0.43*****	**0.22**	**0.38****
	(0.33)	(0.28)	(0.49)		(0.12)	(0.14)	(0.15)
$RCP_t \times MP_t$	**-1.25*****		**-1.62*****	$\Delta RCP_t \times \Delta MP_t$	**-2.42***		**-2.31**
	(0.36)		(0.35)		(1.43)		(1.44)
$RCP_t \times IT_t$		**0.09**	**-0.53**	$\Delta RCP_t \times IT_t$		**1.01*****	**0.74*****
		(0.34)	(0.59)			(0.23)	(0.19)
$RCP_t \times PEG_t$			-1.39**	$\Delta RCP_t \times PEG_t$			-0.23**
			(0.52)				(0.11)
MP_t	0.07***		0.08***	ΔMP_t	-0.003		-0.01
	(0.02)		(0.02)		(0.05)		(0.05)
IT_t		0.11	0.07	IT_t		0.003	0.02
		(0.09)	(0.05)			(0.02)	(0.02)
PEG_t			0.09***	PEG_t			0.03**
			(0.03)				(0.01)
				EC_{t-1}	-0.18***	-0.26***	-0.19***
					(0.03)	(0.03)	(0.03)
				ΔRCP_{t-1}	0.23*	0.32**	0.20*
					(0.12)	(0.15)	(0.10)
				$\Delta REER_{t-1}$	-0.12**	0.05	-0.12**
					(0.05)	(0.05)	(0.05)
R^2	0.47	0.45	0.48		0.21	0.29	0.22
F-statistics	13.28***	8.11***	8.27***		10.95***	46.68***	17.07***
Sample period	1980-2010	1982-2010	1980-2010		1980-2010	1982-2010	1980-2010
# countries	51	23	51		51	23	51
Observations	1,012	644	1,012		1,135	644	1,135

Notes: See notes to Table 3 and 4 for a description of the variable abbreviations and the estimation methods. All specifications in this table include dummy variables to control for 1982-83 Latin American debt crisis (Argentina and Brazil), 1997-98 Asian crisis (Indonesia, Malaysia, Philippines, and Thailand), 2001-02 Argentine crisis, and 2008-09 Global Financial Crisis.

Table 8

Commodity price elasticity and interaction effects: accounting for potential endogeneity.

Estimation method:	Two-step efficient GMM				Two-step system GMM		
Dependent variable:	$REER_t$				$\Delta REER_t$		
	(1)	(2)	(3)		(4)	(5)	(6)
RCP_t	**1.23*****	**1.19*****	**1.30*****	ΔRCP_t	**0.53*****	**0.42*****	**0.41*****
	(0.20)	(0.26)	(0.36)		(0.11)	(0.16)	(0.13)
$RCP_t \times MP_t$	**-1.66*****		**-1.79*****	$\Delta RCP_t \times \Delta MP_t$	**-0.71**		**-3.56**
	(0.44)		(0.46)		(2.31)		(2.67)
$RCP_t \times IT_t$		**0.60***	**0.13**	$\Delta RCP_t \times IT_t$		**1.25*****	**0.90****
		(0.32)	(0.47)			(0.36)	(0.46)
$RCP_t \times PEG_t$			-0.63	$\Delta RCP_t \times PEG_t$			-0.31
			(0.45)				(0.38)
MP_t	0.07**		0.10***	ΔMP_t	-0.01		0.02
	(0.03)		(0.02)		(0.04)		(0.05)
IT_t		0.07	0.05	IT_t		0.03**	0.03**
		(0.07)	(0.04)			(0.01)	(0.01)
PEG_t			0.12***	PEG_t			0.01
			(0.03)				(0.01)
				EC_{t-1}	-0.37***	-0.40***	-0.29***
					(0.04)	(0.04)	(0.03)
				ΔRCP_{t-1}	0.12	0.38***	0.11
					(0.09)	(0.10)	(0.08)
				$\Delta REER_{t-1}$	-0.07*	-0.10	-0.07*
					(0.04)	(0.14)	(0.04)
R^2	0.14	0.13	0.16				
Hansen *J* statistic	0.46	1.25	1.83		47.38	9.53	46.51
Kleibergen-Paap Wald *F* statistic	25.83	15.85	10.76				
AR(1)/AR(2) test (*p*-value)					0.00/0.82	0.08/0.26	0.00/0.91
Sample period	1980-2010	1982-2010	1980-2010		1980-2010	1982-2010	1980-2010
# countries	51	23	51		51	23	51
Observations	1,176	644	1,176		1,135	644	1,135

Notes: See notes to Table 3 and 4 for a description of the variable abbreviations. The regressions in columns (1)-(3) also control for country fixed effects and time trend. Driscoll-Kraay standard errors are reported in parentheses in columns (1)-(3), and Windmeijer standard errors in columns (4)-(6). ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 9
Commodity price elasticity and interaction effects: excluding outliers.

Estimation method:	Dynamic OLS(1,1)				Error correction model		
Dependent variable:	$REER_t$				$\Delta REER_t$		
	(1)	(2)	(3)		(4)	(5)	(6)
RCP_t	**1.11*****	**0.85*****	**1.23*****	ΔRCP_t	**0.45*****	**0.25***	**0.39*****
	(0.21)	(0.20)	(0.32)		(0.10)	(0.13)	(0.13)
$RCP_t \times MP_t$	**-1.14*****		**-1.32*****	$\Delta RCP_t \times \Delta MP_t$	**-1.88**		**-1.71**
	(0.35)		(0.28)		(1.16)		(1.19)
$RCP_t \times IT_t$		**0.75***	**0.11**	$\Delta RCP_t \times IT_t$		**1.11*****	**0.85*****
		(0.37)	(0.40)			(0.29)	(0.22)
$RCP_t \times PEG_t$			-0.94**	$\Delta RCP_t \times PEG_t$			-0.27**
			(0.41)				(0.11)
MP_t	0.08***		0.09***	ΔMP_t	0.004		0.001
	(0.02)		(0.02)		(0.05)		(0.05)
IT_t		0.10	0.08*	IT_t		0.004	0.02
		(0.08)	(0.04)			(0.02)	(0.02)
PEG_t			0.11***	PEG_t			0.04***
			(0.03)				(0.01)
				EC_{t-1}	-0.17***	-0.22***	-0.18***
					(0.03)	(0.05)	(0.03)
				ΔRCP_{t-1}	0.25**	0.35**	0.21**
					(0.11)	(0.15)	(0.10)
				$\Delta REER_{t-1}$	-0.14***	0.06	-0.14***
					(0.05)	(0.06)	(0.04)
R^2	0.48	0.44	0.50		0.17	0.21	0.18
F-statistics	26.09***	13.65***	15.70***		18.50***	31.31***	22.36***
Sample period	1980-2010	1982-2010	1980-2010		1980-2010	1982-2010	1980-2010
# countries	48	22	48		48	22	48
Observations	967	616	967		1,079	616	1,079

Notes: See notes to Table 3 and 4 for a description of the variable abbreviations and the estimation methods. The regressions in this table exclude Ghana, Mozambique, and Paraguay, each of which exhibits extremely large values of long-run commodity price elasticity.

Online Appendix

"Market power, inflation targeting, and commodity currencies"

Yu-chin Chen and Dongwon Lee

Online Appendix A. Elasticity estimates and data characteristics

Table 10
Commodity price elasticity estimates, unit root and cointegration tests for each country.

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)
Argentina		0.11 (1.17)	-1.80 (1)	-1.08 (1)	-0.46 (1)	-0.76 (1)	-2.69 (1)	-4.05	
Australia		3.01*** (0.81)	-1.03 (1)	-1.07 (1)	-0.63 (3)	-0.55 (3)	-2.87 (1)	-4.33	
Bahrain		-0.004 (0.20)	-1.36 (2)	0.88 (1)	-0.54 (3)	-0.87 (1)	-1.56 (1)	-3.95	
Bangladesh	1.47*** (0.31)	-0.23 (0.19)	-2.18 (1)	-1.35 (1)	-1.50 (2)	-0.82 (2)	-3.30* (1)	-4.40*	
Bolivia	0.15 (0.25)	2.36 (1.55)	-0.89 (1)	-0.47 (1)	-0.70 (1)	-0.68 (1)	-2.40 (1)	-6.77***	1985
Brazil		1.02 (1.23)	-1.93 (1)	-1.98 (1)	0.11 (1)	-0.13 (1)	-2.10 (1)	-2.92	
Burundi	3.24*** (0.33)	0.89* (0.44)	-2.14 (1)	-0.65 (1)	-0.83 (1)	-1.01 (1)	-0.80 (1)	-5.16***	2002
Cameroon		-0.60** (0.26)	-2.00 (1)	-1.42 (1)	-0.50 (2)	-0.81 (2)	-1.22 (1)	-3.62	
Canada		1.32*** (0.43)	-1.05 (1)	-1.16 (1)	-0.56 (2)	-0.77 (1)	-1.48 (1)	-3.94	
Central African Rep.	2.07** (0.74)	-0.48 (0.38)	-1.17 (1)	-0.57 (1)	-2.20 (1)	-2.16 (1)	-1.20 (1)	-4.91**	1992
Chile	0.07 (0.41)	0.70** (0.31)	-1.38 (1)	-1.13 (1)	-1.25 (2)	-0.90 (2)	-4.34*** (1)	-3.36	
Colombia		0.85* (0.45)	-1.42 (1)	-1.26 (1)	-0.37 (1)	-0.68 (1)	-1.64 (1)	-3.32	
Costa Rica	0.56 (0.49)	0.11 (0.36)	-0.94 (1)	-1.08 (1)	-0.92 (2)	-0.94 (2)	-10.4*** (0)	-12.56***	1986
Cote d'Ivoire		-0.56 (0.49)	-1.87 (1)	-1.45 (1)	-0.92 (1)	-1.06 (1)	-2.02 (1)	-3.58	
Dominica		0.23 (0.14)	-1.32 (2)	-0.37 (2)	-0.83 (5)	-0.54 (5)	-1.79 (1)	-3.79	
Ethiopia	1.43** (0.68)	0.06 (0.43)	-1.72 (1)	-0.74 (1)	-1.42 (1)	-1.28 (1)	-1.51 (1)	-6.45***	1992
Ghana	7.15** (2.80)	-1.81 (2.01)	-1.22 (1)	0.42 (1)	-0.80 (1)	-0.97 (1)	-2.67 (1)	-4.35*	
Guatemala		0.39 (0.29)	-1.13 (1)	-1.30 (1)	-0.89 (1)	-0.99 (1)	-2.25 (1)	-3.52	
Honduras		-0.26 (0.34)	-1.58 (1)	-1.67 (1)	-1.51 (1)	-1.39 (1)	-1.98 (1)	-3.79	
Iceland		1.79*** (0.63)	-2.55 (1)	-2.18* (1)	-1.36 (1)	-1.03 (1)	-1.90 (1)	-3.38	
India		1.63*** (0.53)	-0.44 (1)	-0.62 (1)	-0.53 (3)	-0.64 (3)	-1.03 (1)	-4.27	
Indonesia		1.73* (0.86)	-0.90 (2)	-0.15 (2)	-0.52 (1)	-0.60 (1)	-1.75 (1)	-3.41	
Kenya	0.31 (0.52)	0.97* (0.48)	-0.97 (1)	-0.84 (1)	-1.16 (1)	-1.09 (1)	-0.53 (1)	-4.41*	
Madagascar	0.83 (0.85)	-0.29 (0.64)	-1.23 (2)	-0.58 (2)	-2.02 (1)	-1.38 (1)	-3.06 (1)	-4.89**	1985
Malawi	0.29 (0.81)	1.65* (0.94)	-0.88 (8)	-0.07 (3)	-0.84 (2)	-0.95 (2)	-0.43 (2)	-4.76**	1992
Malaysia		0.74* (0.40)	-1.96 (1)	-0.38 (1)	-0.60 (1)	-0.74 (1)	-1.61 (1)	-3.35	
Mali		-0.58* (0.33)	-0.72 (1)	-0.88 (7)	0.08 (1)	-0.53 (1)	-0.13 (2)	-3.31	
Mauritania		0.11 (0.29)	-0.79 (1)	0.61 (1)	-0.14 (1)	0.10 (1)	-1.94 (1)	-2.90	
Mauritius		-0.02 (0.10)	-1.73 (1)	-0.68 (1)	-1.26 (2)	-1.06 (2)	-1.86 (1)	-3.61	
Morocco		0.22 (0.19)	-2.17 (1)	-0.49 (1)	-1.30 (3)	-1.13 (3)	-2.57 (1)	-3.87	
Mozambique	10.93** (3.92)	-0.85 (1.35)	-1.18 (1)	0.20 (1)	-1.30 (4)	-1.13 (4)	-2.20 (1)	-4.48*	

Table 10 (continued)
Commodity price elasticity estimates, unit root and cointegration tests for each country.

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)
New Zealand		1.43* (0.80)	-0.76 (6)	0.34 (5)	-0.97 (2)	-0.86 (2)	-2.83 (1)	-3.72	
Niger		-0.001 (0.19)	-0.61 (1)	0.08 (1)	-1.04 (1)	-1.02 (1)	-1.81 (1)	-3.41	
Pakistan		0.55 (0.67)	-0.34 (1)	0.17 (5)	-0.62 (2)	-0.92 (2)	0.01 (1)	-3.19	
Papua New Guinea	1.38*** (0.26)	-0.77 (0.48)	-1.07 (1)	-0.70 (1)	-0.57 (1)	-0.58 (1)	-1.58 (1)	-4.82**	1997
Paraguay	6.20*** (0.71)	1.05* (0.57)	-0.96 (1)	-0.75 (1)	-0.90 (1)	-0.97 (1)	-3.22* (1)	-4.36*	
Peru	-0.65 (0.45)	-1.04 (1.05)	-1.40 (1)	-0.68 (1)	-0.44 (2)	-0.48 (2)	-1.03 (1)	-7.94***	1987
Philippines		0.47 (0.57)	-1.70 (1)	-0.19 (6)	-0.43 (2)	-0.46 (2)	-2.36 (1)	-4.00	
Senegal	0.45 (0.40)	-0.43 (0.42)	-1.62 (1)	-0.53 (1)	-0.69 (3)	-1.07 (3)	-0.42 (1)	-4.97**	1992
South Africa		1.84 (1.13)	-2.55 (1)	-1.02 (1)	-0.05 (1)	-0.28 (3)	-2.43 (1)	-4.13	
Sri Lanka	0.83* (0.43)	0.60** (0.27)	-1.28 (1)	-0.26 (1)	-0.81 (2)	-0.83 (2)	-1.04 (1)	-4.45*	
St. Vincent Gr		0.84*** (0.22)	-1.80 (2)	-0.43 (2)	-0.52 (4)	-0.88 (3)	-1.04 (1)	-4.14	
Sudan		0.20 (0.19)	-1.57 (3)	-1.55 (3)	-0.82 (2)	-0.99 (2)	-3.12 (1)	-3.57	
Suriname		0.47 (0.37)	-1.71 (1)	-2.09 (1)	-0.47 (2)	-0.61 (2)	-2.33 (1)	-2.90	
Tanzania	-0.89 (1.86)	-0.44 (1.24)	-1.88 (1)	-0.45 (1)	-0.09 (1)	-0.69 (1)	-1.23 (1)	-4.86**	1986
Thailand	2.34*** (0.57)	0.69 (0.52)	-1.10 (2)	-0.49 (2)	-0.41 (1)	-0.79 (1)	-0.99 (1)	-4.93**	1997
Togo	1.11*** (0.23)	-0.26 (0.27)	-1.36 (1)	-0.35 (1)	-0.75 (1)	-0.91 (1)	-1.20 (1)	-4.93**	1992
Turkey	2.32** (1.12)	1.29 (1.03)	-0.65 (6)	-1.68 (2)	-0.29 (2)	-0.57 (2)	-1.93 (1)	-4.46*	
Uganda	2.93** (1.23)	0.23 (1.70)	-1.17 (1)	-0.18 (8)	-1.03 (1)	-1.04 (1)	-3.17 (1)	-7.39***	1988
Uruguay		1.18 (0.96)	-2.41 (1)	-2.33* (1)	-0.68 (4)	-1.31 (4)	-2.20 (1)	-3.35	
Zambia		0.41 (0.83)	-0.80 (3)	-1.69 (3)	-1.04 (3)	-0.93 (2)	-2.64 (1)	-3.22	

Notes: Columns (1) and (2) present commodity price elasticity estimates with Newey-West HAC standard errors in brackets. The long-run commodity price elasticity of the real exchange rate is estimated by the dynamic OLS ($REER_t = \beta_0 + \beta_1 RCP_t + \sum_{j=-1}^{1} \gamma_j \Delta RCP_{t+j} + u_t$), and the short-run elasticity by the first differencing ($\Delta REER_t = \delta_0 + \delta_1 \Delta RCP_t + \epsilon_t$), where both RCP and $REER$ are measured in logarithm. Columns (3)-(6) report test statistics of DF-GLS unit-root test (Elliot et al., 1996) for $REER$ and RCP with and without a deterministic trend term. The lag length is automatically chosen due to the minimum of the modified Akaike information criterion and presented in parentheses. Column (7) documents the Augmented Engle-Granger (AEG) cointegration test statistics and their level of significance (based on the critical values from MacKinnon (2010)) with the number of optimal lags chosen by the Schwarz Bayesian information criterion reported in parentheses. Columns (8) and (9) show the Gregory-Hansen (1996) test statistics and associated structural shift dates. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 11
Cross sectional dependence, panel unit root and cointegration tests.

a. Cross-sectional dependence test

	(1)	(2)
CD test statistic	44.8***	13.3***
Average absolute value of the off-diagonal elements	0.5	0.2

b. Panel unit-root tests

		$REER$	RCP	MP	TO	FO	CEX	RES
LLC	Levels	0.8	4.0	-3.1***	1.7	-3.1***	-0.2	3.2
	1st diff.	-13.3***	-2.2**	-27.1***	-23.4***	-27.3***	-18.7***	-21.1***
IPS	Levels	1.8	9.2	-2.5***	0.75	-2.4***	1.7	2.0
	1st diff.	-16.4***	-4.5***	-25.5***	-20.4***	-25.0***	-16.8***	-20.6***
CIPS	Levels	-2.0**	-4.4***	-3.6***	1.3	-1.3	-0.9	-1.7**
	1st diff.	-17.1***	-20.5***	-12.0***	-13.5***	-11.8***	-13.2***	-14.0***

c. Panel cointegration tests

Kao			
ADF t-statistic	-4.3***		
Pedroni			
Within-dimension		*Between-dimension*	
Panel v-statistic	1.0	Group ρ-statistic	-2.1**
Panel ρ-statistic	-3.6***	Group PP-statistic	-9.2***
Panel PP-statistic	-8.3***	Group ADF-statistic	-5.2***
Panel ADF-statistic	-4.5***		
Westerlund			
G_t	-3.0***		
G_a	-12.6		
P_t	-21.6***		
P_a	-12.5***		

Notes: $REER$ = real effective exchange rate; RCP = real commodity price; MP = market power; TO = trade openness; FO = financial openness; CEX = commodity export dependency; RES = international reserves. In panel a, Pesaran (2004)'s cross-sectional dependence (CD) test statistic is based on the residuals of the regression model specifications: (1) $REER_{it} = c_i + \beta_1 RCP_{it} + u_{it}$; and (2) $\Delta REER_{it} = c_i + \delta_1 \Delta RCP_{it} + \epsilon_{it}$. 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. Reported are the t^*-statistic for the LLC test (Levin et al., 2002), W-statistic for the IPS test (Im et al., 2003), and $Z[t\text{-bar}]$ statistic for the CIPS test (Pesaran, 2007). In panel c, for the Kao (1999) test, an individual intercept is included only, while the individual intercept and individual trend are included for the Pedroni (2004) test. For the Westerlund (2007) test, we set the width of Bartlett-kernel window at 3 and allow for a constant and deterministic trend in the cointegrating relationship. ***, ** 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% 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 which uses AIC.

Table 12
Panel Granger non-causality test.

Null hypothesis	$\tilde{Z}$ statistic	p-value
ΔRCP does not Granger-cause $\Delta REER$.	133.74	< 0.01
$\Delta REER$ does not Granger-cause ΔRCP.	0.51	0.61

Notes: The alternative hypothesis is that ΔRCP ($\Delta REER$) does Granger-cause $\Delta REER$ (ΔRCP) for at least one country. $\tilde{Z}$ statistic is the cross-section average of individual Wald statistics, which will asymptotically converge in distribution to a normal distribution. The optimal number of lags of the series is selected based on the Akaike information criterion.

Online Appendix B. Choice of sample countries and commodities

We select our sample by keeping commodity-dependent developing countries whose export earnings in nonfuel primary products accounted for more than half of total export earnings for the years 1988-92.[1] From 73 countries based on this classification, 20 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. (2011), Ecuador was excluded because of its dollarization that began in 2001, and Nicaragua because of the unusual 1000% appreciation at the beginning of the sample period. Zimbabwe was dropped as well due to the hyperinflation during the significant part of sample period since 2002. In addition, Mexico, Syrian Arab Republic, and Tunisia were removed because of their heavy dependence on the crude oil whose average export share in aggregate commodity exports was greater than 50% over the sample period. Lastly, we added four commodity-dependent developed countries (Australia, Canada, Iceland, and New Zealand) to our sample. This procedure leaves a total of 51 commodity-producing countries. The full list of countries is available in Table 10 or 14.

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

Table 13
List of commodities employed in the construction of commodity price indices.

Type	Commodity
Agricultural food	Bananas, Barley, Beef, Cocoa, Coconut oil, Coffee, Copra, Fish, Fishmeal, Groundnuts (peanuts), Groundnut oil, Lamb, Maize, Olive oil, Oranges, Palm kernel oil, Palm oil, Poultry (chicken), Rapeseed oil, Rice, Shrimp, Sorghum, Soybean meal, Soybean oil, Soybeans, Sugar, Sunflower oil, Swine, Tea, Tobacco, Wheat
Agricultural raw materials	Cotton, Hard logs, Hardwood sawn, Hides, Rubber, Soft logs, Softwood sawn, Wood pulp, Wool (coarse), Wool (fine)
Fertilizers	Phosphate rock, Potash, TSP (triple superphosphate), Urea
Metals	Aluminum, Copper, Gold, Iron ore, Lead, Nickel, Silver, Tin, Uranium, Zinc
Non-oil energy	Coal, Natural gas
Oil	Crude oil (petroleum)

Table 14
Primary 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
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	Hardwood sawn	0.48	0.13	0.10	0.07	0.06
Canada	Crude oil	Natural gas	Softwood sawn	Wood pulp	Wheat	0.17	0.16	0.12	0.10	0.07
Central African Rep.	Hard logs	Cotton	Hardwood sawn	Coffee	Soft logs	0.38	0.32	0.17	0.15	0.05
Chile	Copper	Natural gas	Wood pulp	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	Hardwood sawn	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	Hardwood sawn	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
Madagascar	Shrimp	Coffee	Sugar	Cocoa	Hardwood sawn	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	
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

Table 14 (continued)
Primary 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
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
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
Senegal	Fish	Groundnut oil	Phosphate rock	Crude oil	Cotton	0.26	0.24	0.19	0.09	0.09
South Africa	Coal	Aluminum	Iron	Wood pulp	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
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
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
Uruguay	Beef	Rice	Fish	Wool (coarse)	Soybeans	0.41	0.20	0.11	0.07	0.05
Zambia	Copper	Sugar	Cotton	Tobacco	Maize	0.86	0.04	0.03	0.03	0.01

Notes: Reported are top five major commodities exported by each country for the period 1980-2010. 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. Source: Calculations in this table are based solely on the data from UN COMTRADE.

Online Appendix C. Model selection: least angle regressions

As a way to verify our baseline empirical specifications in Eqs. (7) and (8), we implement a model-selection algorithm. We use least angle regressions (LARS) of Efron et al. (2004) that consider parsimony as well as prediction accuracy. The LARS procedure provides a natural way to judge the relative importance of the variables explaining $REER$ and $\Delta REER$, and is superior to the traditional stepwise regression.[2] LARS begins by setting the coefficients on all predictors to zero, and adds in variables step-by-step based on their correlation with the residuals of the previous model. To select the shrinkage level (the number of variables to include), we use the minimized C_p criterion, where C_p is an estimate of the prediction error. The goal is to see if variables of our interest would be selected for inclusion in the specifications producing the minimum C_p.

Table 15 column (1) presents the long-run DOLS(1,1) specification chosen by the minimized C_p statistics and the coefficient estimates for the chosen variables. In addition, the cumulative R^2's for the regression after the inclusion of the particular variable are reported in the parentheses. For example, we see that in the long-run regression with all possible elasticity determinants (the specification employed in section 4.3.1), the first variable selected is CEX, since it has the smallest R^2 amongst the reported numbers. The next variable entering is the IT interaction term ($RCP \times IT$), followed by TO interaction term ($RCP \times TO$), CEX interaction term ($RCP \times CEX$) and so on, producing a final R^2 of 0.5601. Note that all key variables emphasized in this and the other studies are selected under LARS from the long-run model.[3]

[2] See Madigan and Ridgeway (2004) for more details.

[3] Other variables are also chosen under LARS such as contemporaneous, 1 lead and 1 lag of changes of RCP and interaction terms. However, for brevity, we list coefficients for selected variables of our main interest and their goodness-of-fit measures in Table 15.

The LARS procedure is also applied to the short-run model and the results are reported in Table 15 column (2). The algorithm selects ΔMP interaction term ($\Delta RCP \times \Delta MP$), IT interaction term ($\Delta RCP \times IT$), contemporaneous and lagged changes of RCP, lagged differences of $REER$, and lagged error correction term from the ECM specification. We also note a generally weaker short-run result in terms of the final R^2.

Overall, the LARS model selection method strongly supports our empirical model specifications employed to estimate the long-run and short-run elasticities and their determinants.

Table 15

Least angle regressions: coefficient values based on minimum C_p statistics.

Long-run model: DOLS(1,1)		Short-run model: ECM	
Dependent variable: $REER_t$		Dependent variable: $\Delta REER_t$	
	Coefficient estimate (R^2)		Coefficient estimate (R^2)
	(1)		(2)
RCP_t	1.1402 (0.5601)	ΔRCP_t	0.2783 (0.0806)
$RCP_t \times MP_t$	-1.2689 (0.4769)	EC_{t-1}	-0.1345 (0.0453)
$RCP_t \times IT_t$	-0.0823 (0.2015)	$\Delta RCP_t \times \Delta MP_t$	-0.1273 (0.1732)
$RCP_t \times TO_t$	-5.5984 (0.2528)	$\Delta RCP_t \times IT_t$	0.6247 (0.1267)
$RCP_t \times FO_t$	1.0617 (0.4855)	$\Delta RCP_t \times \Delta TO_t$	0.7978 (0.1559)
$RCP_t \times CEX_t$	0.5803 (0.3261)	$\Delta RCP_t \times \Delta RES_t$	-1.9997 (0.1623)
$RCP_t \times RES_t$	4.4697 (0.4786)	ΔMP_t	0.0140 (0.1722)
$RCP_t \times PEG_t$	-0.9007 (0.4982)	IT_t	0.0060 (0.1607)
MP_t	-0.0019 (0.4569)	ΔTO_t	-0.3161 (0.0551)
IT_t	0.0349 (0.5223)	ΔRCP_{t-1}	0.2044 (0.1484)
TO_t	-0.2977 (0.4028)	$\Delta REER_{t-1}$	-0.1316 (0.0751)
FO_t	0.3196 (0.3738)		
CEX_t	0.3551 (0.0841)		
RES_t	0.4858 (0.4883)		
PEG_t	0.0975 (0.3502)		

Notes: See notes to Table 3 and 4 for a description of the variable abbreviations and the estimation methods. Reported are coefficient estimates for regressors chosen based on the minimum C_p statistic under least angle regressions. Numbers in the parentheses represent the total R^2 when the particular regressor is added.

Online Appendix D. Heterogeneous international reserve accumulation patterns

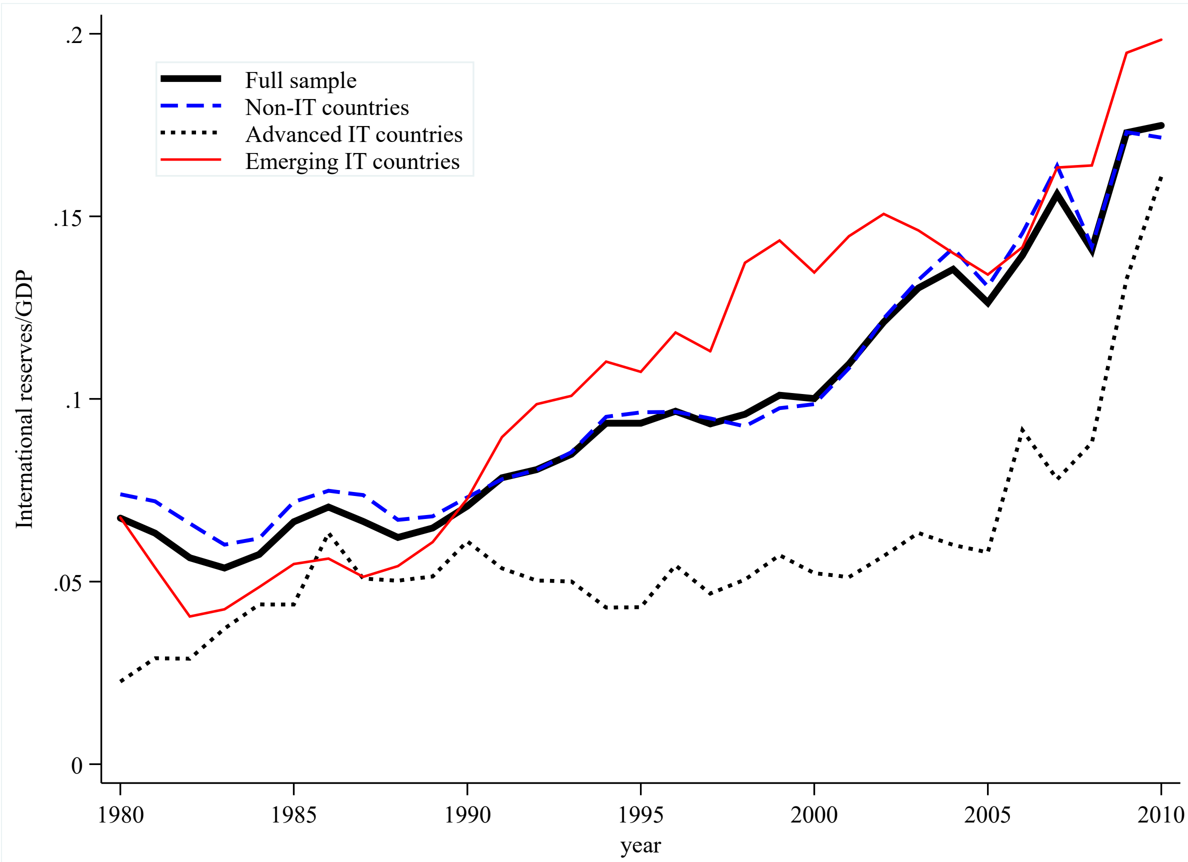

Fig. 3. Average time series patterns of international reserves by IT (inflation targeting) and non-IT countries.

Online Appendix E. Additional robustness test results

Table 16
Endogeneity test for commodity prices.

Estimation method:	Two-step efficient GMM		
	a. *First-stage regression*		
Dependent variable:	RCP_t		
	(1)	(2)	(3)
RCP_{t-2}	0.72***		0.47***
	(0.08)		(0.11)
RCP_{t-3}		0.68***	0.32**
		(0.10)	(0.14)
R^2	0.44	0.40	0.47
Kleibergen-Paap rank Wald *F* statistic	85.10	50.99	40.62
	b. *Second-stage regression*		
Dependent variable:	$REER_t$		
	(1)	(2)	(3)
RCP_t	1.44***	1.33***	1.33***
	(0.23)	(0.21)	(0.21)
R^2	0.21	0.19	0.19
p-value			
Hansen *J* statistic			0.99
Endogeneity test: Hansen *C* statistic	**0.07**	**0.09**	**0.08**
# countries	51	51	51
Observations	1,479	1,428	1,428

Notes: $REER$ = real effective exchange rate; RCP = real commodity price. The regressions also control for country fixed effects and time trend. Driscoll-Kraay standard errors are reported in parentheses. ***, ** indicate statistical significance at the 1% and 5% levels, respectively.

Table 17
Short-run commodity price elasticity and interaction effects: Hausman-Taylor estimator.

Estimation method:	Hausman-Taylor GLS					
Dependent variable:	$\Delta REER_t$					
	(1)	(2)	(3)	(4)	(5)	(6)
ΔRCP_t	**0.38*****	**0.44*****	**0.24****	**0.24****	**0.38*****	**0.40*****
	(0.08)	(0.10)	(0.10)	(0.11)	(0.09)	(0.12)
$\Delta RCP_t \times \Delta MP_t$		**-2.00***		**-0.49**		**-1.83***
		(1.07)		(1.14)		(1.08)
$\Delta RCP_t \times IT_t$			**1.06*****	**1.06*****	**0.83*****	**0.79*****
			(0.33)	(0.33)	(0.29)	(0.30)
$\Delta RCP_t \times PEG_t$					-0.31**	-0.27
					(0.14)	(0.19)
ΔMP_t		0.005		-0.04		0.002
		(0.04)		(0.03)		(0.04)
IT_t			0.01	0.01	0.03***	0.02**
			(0.01)	(0.01)	(0.01)	(0.01)
PEG_t					0.05*	0.04
					(0.03)	(0.03)
EC_{t-1}	-0.20***	-0.17***	-0.24***	-0.23***	-0.21***	-0.19***
	(0.02)	(0.03)	(0.04)	(0.06)	(0.02)	(0.03)
ΔRCP_{t-1}	0.28***	0.19**	0.31***	0.16	0.25***	0.15**
	(0.08)	(0.08)	(0.11)	(0.10)	(0.09)	(0.08)
$\Delta REER_{t-1}$	-0.02	-0.09	0.06	-0.05	-0.02	-0.09
	(0.05)	(0.06)	(0.06)	(0.06)	(0.05)	(0.06)
R^2	0.14	0.14	0.25	0.24	0.14	0.14
Chi-square statistics		33.63***	21.21***	18.77***	32.70***	36.16***
Sample period	1980-2010	1980-2010	1982-2010	1982-2010	1980-2010	1980-2010
# countries	51	51	23	23	51	51
Observations	1,479	1,135	644	515	1,479	1,135

Notes: See notes to Table 4 for a description of the variable abbreviations. Robust standard errors are reported in parentheses. Chi-square statistics for a Wald test and their significance level are reported to test the joint significance of coefficients for ΔRCP_t and interaction terms. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 18
Short-run commodity price elasticity and interaction effects: allowing for panel heterogeneity.

Estimation method:	Pooled mean group estimator							
Dependent variable:	$\Delta REER_t$							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Long-run coefficient								
RCP_t	**1.26*****	**0.65****	**1.01*****	**2.20*****	**2.72*****	**0.53****	**1.79*****	**2.37*****
	(0.25)	(0.27)	(0.16)	(0.11)	(0.12)	(0.27)	(0.15)	(0.08)
Short-run coefficients								
ΔRCP_t	**0.46****	**0.39**	**-0.05**	**0.19**	**0.34**	**0.22**	**-0.37**	**0.05**
	(0.23)	(0.40)	(0.25)	(0.32)	(0.22)	(0.32)	(0.24)	(0.28)
$\Delta RCP_t \times \Delta MP_t$	**-7.69**			**11.77**	**-28.17**			**-40.37**
	(26.02)			(24.31)	(17.12)			(25.21)
$\Delta RCP_t \times IT_t$		**0.08**	**0.31**	**0.48***		**0.31**	**0.47****	**0.51****
		(0.32)	(0.23)	(0.26)		(0.31)	(0.22)	(0.24)
$\Delta RCP_t \times PEG_t$			0.87**	0.43			0.80**	0.39
			(0.43)	(0.35)			(0.37)	(0.29)
ΔMP_t	0.46			0.16	-0.33			0.16
	(1.51)			(1.59)	(0.77)			(0.76)
IT_t		-0.04**	-0.01	-0.02**		-0.04**	-0.01	-0.02**
		(0.02)	(0.01)	(0.01)		(0.02)	(0.004)	(0.01)
PEG_t			0.02	0.003			0.02	0.003
			(0.02)	(0.01)			(0.01)	(0.01)
EC_{t-1}	-0.01***	-0.29***	-0.26***	-0.03***	-0.03**	-0.27***	-0.28***	-0.03**
	(0.005)	(0.04)	(0.03)	(0.01)	(0.01)	(0.04)	(0.03)	(0.01)
ΔRCP_{t-1}	0.11	0.13	0.08	0.19	0.07	-0.09	-0.18	0.07
	(0.16)	(0.23)	(0.12)	(0.16)	(0.13)	(0.24)	(0.14)	(0.15)
$\Delta REER_{t-1}$	-0.13***	0.01	-0.01	-0.15***	-0.11**	0.01	0.0002	-0.15***
	(0.03)	(0.05)	(0.03)	(0.04)	(0.05)	(0.04)	(0.03)	(0.05)
Cross-sectional dependence control	No	No	No	No	Yes	Yes	Yes	Yes
Log-likelihood	1255.87	592.80	1413.92	1310.42	1366.43	640.12	1511.93	1423.44
Sample period	1980-2010	1982-2010	1980-2010	1980-2010	1980-2010	1982-2010	1980-2010	1980-2010
# countries	51	23	51	51	51	23	51	51
Observations	1135	621	1479	1135	1135	621	1479	1135
Obs per country (avg)	22.3	27	29	22.3	22.3	27	29	22.3

Notes: See notes to Table 4 for a description of the variable abbreviations. The regressions in columns (5)-(8), unlike columns (1)-(4), include the cross-section averages of changes in $REER$ as an additional regressor to account for cross-sectional dependence in residuals. All specifications include country fixed effects. Standard errors are reported in parentheses. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

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