Commodity currency reactions and the Dutch disease:

The role of capital controls

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Abstract

Commodity windfall gains generally induce real exchange appreciations in commodity-rich economies and make other tradable sectors less competitive in global markets. This Dutch disease phenomenon has been blamed for causing slow growth. Based on the theory, we hypothesize that applying capital controls may mitigate the transmission of positive commodity price shocks to the real exchange rate and help shield manufactured exports. Examining a panel dataset of 37 developing countries over the period from 1980 to 2017, we find that a more excessive commodity currency appreciation indeed has a more detrimental impact on the export performance of the manufacturing sector. Restrictions on capital inflows tend to curb real appreciation pressures and alleviate the severity of the Dutch disease in accordance with our hypothesis. Our findings suggest the countercyclical use of capital controls in commodity-exporting countries to foster economic diversification and improve their growth potential.

Keywords: Capital controls; Commodity price; Dutch disease; Manufactured exports; Real exchange rate

JEL classification: F31; F32; O13; Q33

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1. Introduction

Commodity-rich economies often face large fluctuations in the value of their currencies due to volatile global prices of their primary exports. These currency fluctuations can have detrimental impacts on the local economy. For example, persistent real appreciations could lead to reduced competitiveness and investment in non-commodity export sectors (e.g., manufacturing). Conversely, sharp depreciations could increase the debt burden on domestic firms with large foreign liabilities. For these reasons, maintaining a competitive and stable exchange rate may be of special interest to commodity-abundant developing countries pursuing economic diversification and an export-led growth strategy.

In this paper, we focus on the effectiveness of capital controls in stabilizing the real exchange rate and preserving the competitiveness of manufactured exports in commodity-dependent developing economies. Since the manufacturing sector is known for its positive externalities in production, such as learning-by-doing and knowledge spillovers (van Wijnbergen, 1984; Krugman, 1987; Matsuyama, 1992; Sachs and Warner, 1995; Gylfason et al., 1999; Torvik, 2001), our result has the potential to help design sustained growth policies in developing countries susceptible to the Dutch disease.[1]

To understand the importance of manufactured exports to economic development, Fig. 1 displays relevant historical evidence in our sample of developing countries over the past three-and-a-half decades. In the figure, each country has two observations for the log of manufactured

[1] The Dutch disease refers to the coexistence of booming and lagging tradable goods sectors in a resource-rich economy that generally suffers from low economic growth despite its large endowment of raw commodities. The disease can arise from various forms of shocks such as a large natural resource discovery, a rise in the commodity price, or large inflows of foreign aid or remittances. For seminal articles in this topic, see Corden and Neary (1982) and Corden (1984) for theoretical developments and Sachs and Warner (1995, 1999, 2001) for supporting empirical evidence. Also, see Frankel (2010), van der Ploeg (2011), and Magud and Sosa (2013) for an extensive review of the literature, and Harding and Venables (2016) for a recent empirical exploration.

exports (as a ratio to GDP) and the growth rate of per capita GDP, which are averages for each period, 1980−1999 and 2000−2017, so that we can trace their temporal changes within the economy. From the illustration, we detect an apparent positive relationship between those two variables in our sample when other standard growth determinants are also controlled. In line with this observation, Hausmann et al. (2007), Jones and Olken (2008), Johnson et al. (2010), Berg et al. (2012), and Sheridan (2014) argue that growth accelerations are strongly associated with expansions of the manufacturing export sector.[2]

[Insert Fig. 1 here]

The result in Fig. 1 suggests that developing countries with heavy reliance on primary commodity products may have an incentive to diversify their economies with an expansion of the manufacturing sector that provides momentum for long-run economic growth. The main purpose of this paper is to explore how those countries may achieve such a development objective by managing their capital accounts and real exchange rate behavior.

Building on a simple static open macroeconomic model by Obstfeld and Rogoff (1996), we first present the theoretical underpinnings of the economic structure in a commodity-abundant country that is assumed to produce exportable commodities and manufactured goods as well as nontraded goods. In such an economy, a rise in the world price of the country's commodity exports tends to appreciate its real exchange rate, whose reaction magnitude depends on the degree of capital account openness.

The theoretical framework generates two testable hypotheses: First, capital controls

[2] In a related vein, Dabla-Norris et al. (2010) show that the impact of foreign direct investment (FDI) on economic growth is significantly positive only for countries with more diversified economic structures (i.e., lower dependence on commodity exports).

mitigate the transmission of commodity price shocks to the real exchange rate. Second, capital controls reduce the propensity to crowd out manufactured exports resulting from a commodity price boom.

To explicitly test these hypotheses, we undertake a systematic panel data analysis based on a sample of 37 non-oil commodity-exporting developing countries over the years from 1980 to 2017. Using the export volumes of 58 primary commodities and their global prices, we first construct a country-specific real commodity price index as in Cashin et al. (2004) and Chen and Lee (2018). We then show that commodity prices and real exchange rates are cointegrated and exhibit a strong long-run comovement in our sample countries. In addition, we find statistically significant evidence that capital controls, especially on FDI inflows (most likely toward the commodity industries), help avoid a sharp real appreciation following a surge in commodity prices.

Recognizing commodity prices as a driving force in the evolution of real exchange rates, we find that capital account restrictions tend to shield manufactured exports by reducing the real appreciation pressures stemming from a steep increase in commodity prices. In support of capital controls' positive role of preserving export competitiveness, we also report that the more excessive the commodity currency appreciation or real overvaluation, the worse the export performance of manufacturing. These results suggest the countercyclical use of capital controls in countries whose currency values are strongly tied to their commodity export prices to lower the intensity of the Dutch disease.

Our baseline results are robust to using alternative measures of capital controls based on de jure and hybrid financial openness indices and controlling for exchange rate regimes and major financial crises.

This paper contributes to a vast literature on the Dutch disease and real exchange rate in the following three ways. First, we disentangle the dynamics of the Dutch disease into two key links—one from commodity prices to real exchange rates and the other from the real exchange rates to manufactured exports—and jointly address them in this paper. These relationships have typically been studied separately in the prior literature. For example, the first link has been analyzed in the commodity currency literature, such as Amano and van Norden (1995), Chen and Rogoff (2003), Cashin et al. (2004), Coudert et al. (2011), Ricci et al. (2013), Bodart et al. (2012, 2015), and Chen and Lee (2018). The second link has been investigated by Grobar (1993), Sekkat and Varoudakis (2000), Prasad et al. (2007), and Rajan and Subramanian (2011), who report the damaging influence of real exchange rate uncertainty or misalignment. Unlike these studies, we examine the impacts of commodity price shocks on manufactured exports, with the degree of real exchange rate reaction determining the severity of the Dutch disease.

Second, our findings enrich the debate in the literature regarding how effective capital controls are at managing unfavorable real exchange rate movements. Bodart et al. (2015) find that, contrary to our results, an increase in commodity prices is related to stronger real appreciation when a country has a less open capital account. Magud et al. (2018) survey close to 40 empirical studies and conclude that capital controls may help retain monetary autonomy and alter the composition of capital flows; however, there are only a few successful cases in reducing real appreciation pressures: in Chile, Malaysia, and Thailand. By contrast, Erten and Ocampo (2016) find that capital account regulations are useful to decrease a real appreciation in emerging economies. Similarly, some studies find that developing countries with higher capital account openness are more likely to experience real overvaluation (Prasad et al., 2007) or less undervaluation (Rodrik, 2008). The present paper complements this last strand of the literature.

Relative to the prior work, however, we emphasize the role of capital controls in limiting the transmission of commodity price changes into the real exchange rate, a particularly relevant concern for commodity-rich developing countries.

Third, we attempt to extend the Dutch disease literature using a sample of non-oil commodity exporters and their export price movements as a source of foreign exchange windfall shocks.[3] Using such external shocks provides clear identification advantages in the empirical models of the Dutch disease. This can be justified by the notion that the world commodity price dynamics are driven mostly by global supply and demand conditions and can serve as an exogenous terms-of-trade shock to the vast majority of commodity exporters (Chen et al., 2010). As such, in contrast to the regression models that address a link between remittances and the real exchange rate (e.g., Amuedo-Dorantes and Pozo, 2004; Lartey et al., 2012) or foreign aid flows and economic growth (e.g., Rajan and Subramanian, 2005, 2008), it is less likely that our models suffer from a potential endogeneity bias.[4]

As is widely known, international financial integration can offer various macroeconomic benefits. For example, portfolio equity or debt inflows can relieve the financing constraints of developing countries that otherwise face the high cost of capital with limited borrowing sources. FDI inflows can bring along state-of-the-art technologies and managerial skills and improve market accessibility. Growing financial integration also increases diversification opportunities for both domestic and foreign investors.

Nevertheless, our findings indicate that countercyclical capital controls appear to be a desirable policy toolkit in commodity-exporting developing countries to effectively manage real

[3] Similar to our work, Ismail (2010) evaluates the Dutch disease effects of permanent oil price shocks using a small set of oil-exporting countries.

[4] In the earlier literature, reverse causality was a potential concern because "migrants usually look at exchange rates in order to decide how much to remit back home" (Lartey et al., 2012); and "aid flows could go to countries that are doing particularly badly, or to countries that are doing well" (Rajan and Subramanian, 2008).

appreciation pressures arising from their export price booms and to implicitly subsidize economic diversification.[5,6] In line with this view, Aizenman et al. (2007) and Prasad et al. (2007) insist that higher ratios of self-financing may spur faster growth when nonindustrial countries do not have adequate capacity to absorb foreign resources due to unstable macroeconomic policies and economic structures that are vulnerable to overvaluations.

In the next section, we present a simple small open economy model and derive two testable hypotheses. Section 3 describes the data and empirical model specifications. The baseline estimation results and robustness analyses are reported in Section 4, and finally, Section 5 concludes.

2. A theoretical framework

This section presents a three-sector small open economy model that highlights the transmission of commodity price shocks to the real exchange rate and the resulting response in exports of manufactured goods. The model builds on the canonical framework of Obstfeld and Rogoff (1996, Ch. 4), with relevant implications taken from Corden and Neary (1982) and Bodart et al. (2015).

For our purposes, we assume that all of the commodity goods produced by the home country are exported abroad, but the foreign country in the model is not involved in commodity

[5] In fact, imposing capital controls can avoid selecting beneficiaries for export subsidies and uniformly provide an economy-wide incentive to all exporting industries.

[6] For capital controls and their role as a macroprudential policy, see the recent surveys provided in Engel (2016) and Erten et al. (forthcoming).

trading at all. As is standard in the literature, we let global commodity prices be determined by the world market conditions and thus exogenously be given to the domestic commodity sector.

The parsimonious model structure enables us to derive three propositions, which form the basis of our main hypotheses. The detailed model derivations can be found in Appendix E.

2.1. Production

Consider that the domestic economy produces three types of goods: exportable commodities or resources (R), exportable manufactured goods (M), and nontraded (N) goods. The production function in each sector exhibits constant returns to scale and is given by

$$Y_R = A_R L_R^{\alpha} K_R^{1-\alpha}, \tag{1}$$

$$Y_M = A_M L_M^{\beta} K_M^{1-\beta}, \tag{2}$$

$$Y_N = A_N L_N, \tag{3}$$

where A_i, L_i, and K_i are the total factor productivity, labor, and capital stock employed in the production of sector $i = R, M, N,$ respectively. Note that both capital and labor are required in the production of tradable goods, with α and β capturing the labor share. The nontraded goods' production is assumed to rely on labor as the only input.

In the benchmark case, we assume that capital is perfectly mobile internationally and labor is mobile only domestically. Thus, the domestic marginal product of capital is given by the world interest rate r^*, while perfect domestic labor mobility ensures that the wage rate w is equalized across sectors. Like Obstfeld and Rogoff (1996), we allow a common rate of productivity shocks in the exportable sectors.

Under the assumptions above, combining log-differentiated profit-maximization conditions in three sectors gives

$$\hat{p}_N = \tau(\hat{p}_R - \hat{p}_M) - \hat{A}_N \tag{4}$$

with $\tau = 1/(\mu_{L,R} - \mu_{L,M})$,

where p_i is the price of goods in sector i; $\mu_{L,i}$ is the labor income share $\left(0 < \mu_{L,i} < 1\right)$, defined as $\mu_{L,i} \equiv wL_i/p_iY_i$; and a hat above the variable denotes a logarithmic derivative, $\hat{x} = \mathrm{d}(\ln x)$. Note that as long as the commodity sector is more labor-intensive than manufacturing, we have $\tau > 0$.[7] The underlying mechanism in Eq. (4) is that higher commodity prices (relative to prices in the manufacturing sector) raise the demand for labor and the wage rate in the commodity sector. This in turn causes a shift of labor out of the other sectors and an increase in the overall wage rate, eventually boosting the price of labor-intensive nontraded goods.

2.2. Consumption

The representative domestic household in our model economy consumes two types of goods: nontraded and manufactured products. Accordingly, a domestic consumer's utility function takes the following Cobb-Douglas form:

[7] Equivalently, the manufacturing sector is assumed to be more capital-intensive than the commodity sector. This assumption is needed to replicate the main logic of the resource movement effect that follows (as in Corden and Neary, 1982).

$$U = \gamma C_N^{\theta} C_M^{1-\theta}, \tag{5}$$

where C_N and C_M are the consumption of the two goods, θ is the share of nontraded goods in the domestic household's consumption basket, and $\gamma = \theta^{-\theta}(1-\theta)^{-(1-\theta)}$.

Similarly, the representative household in the foreign country consumes the nontraded goods as well as imported manufactured goods that are produced by the home country. These two goods are not perfect substitutes for foreign consumers. A foreign household shows the following preferences:

$$U^* = \gamma^* C_N^{*\theta^*} C_M^{*1-\theta^*}, \tag{6}$$

where $\gamma^* = \theta^{*-\theta^*}(1-\theta^*)^{-(1-\theta^*)}$ and a superscript asterisk on the variable denotes a foreign value.

Note that since the supply of nontraded goods satisfies the domestic demand and the labor supply is fixed in the domestic factor market, the market clearing conditions in the home country are given by

$$Y_N = C_N, \tag{7}$$

$$L = L_R + L_M + L_N. \tag{8}$$

2.3. Real exchange rate

In the absence of any frictions in international trade, the law of one price is assumed to hold in the long run for the tradable goods so that

$$Ep_i = p_i^* \text{ for } i = M, R, \tag{9}$$

where E is the nominal exchange rate, defined as the price of domestic currency in terms of foreign currency, and p_i and p_i^* are the domestic and foreign currency prices of tradable good i, respectively.

Using the consumption-based price index for the home and foreign economies and the law of one price for tradable goods, we can express the real exchange rate (Q), the relative price of the domestic consumption basket in terms of the foreign consumption basket, as follows:

$$Q = \frac{EP}{P^*} = \frac{Ep_N^{\theta} p_M^{1-\theta}}{(p_N^*)^{\theta^*}(p_M^*)^{1-\theta^*}}, \tag{10}$$

where P and P^* are domestic and foreign aggregate price indices, respectively. By construction, an increase in Q indicates a real appreciation of the home currency relative to the foreign currency.

2.4. Model implications

This subsection summarizes three propositions that emerge from the model.

Proposition 1. *An increase in world prices of commodities induces real appreciation in a commodity-exporting country.*

Proof. By log-differentiating Eq. (10) and combining the result with Eq. (4) and the law of one price for the tradable goods, we can find the following marginal effect of a positive shock in global commodity prices on the real exchange rate:

$$\frac{\partial \hat{Q}}{\partial \hat{p}_R^*} = \tau\theta > 0. \tag{11}$$

Given that labor is perfectly mobile between sectors and the price in the manufacturing sector is internationally determined, the higher demand for labor in the commodity sector following a surge in commodity prices raises the overall wage rate. This in turn bids up the prices of nontraded goods and gives rise to a real exchange rate appreciation.[8] ■

Proposition 2. *A commodity price boom crowds out manufactured exports through the real appreciation.*

Proof. To simplify the matter, let exports and imports of manufactured products rely on their relative prices:[9]

$$X_M = X_M\left(\frac{p_M}{P}\right), \tag{12}$$

$$C_M^* = C_M^*\left(\frac{p_M^*}{P^*}\right), \tag{13}$$

[8] According to Eq. (11), the larger the size of parameter θ (the share of nontraded goods in domestic consumption), the larger the real exchange rate response to an increase in commodity prices. This arises because price changes in commodity exports are transmitted into the real exchange rate primarily through adjustments in nontraded good prices. For related discussions and supporting empirical evidence, see Bodart et al. (2015) and Chen and Lee (2018).

[9] Clements and Fry (2008) use a similar analytical framework to describe the equilibrium in the world commodity market.

where the definition of manufactured exports is given by subtracting domestic consumption from production such that $X_M \equiv Y_M - C_M$. Since the two countries, home and foreign, determine market forces, the world market clears when $X_M = C_M^*$. By log-differentiating this market-clearing condition, combined with the law of one price for tradable manufacturing sector and the definition of the real exchange rate, we find

$$\widehat{\left(\frac{p_M^*}{P^*}\right)} = \eta \hat{Q}, \tag{14}$$

where $\eta = \varepsilon^s / (\varepsilon^s - \varepsilon^d)$, $\varepsilon^s (\geq 0)$ is the price elasticity of manufacturing supply, and $\varepsilon^d (\leq 0)$ is the price elasticity of manufacturing demand. Since $0 \leq \eta \leq 1$, Eq. (14) shows a positive relationship between the foreign relative price of manufactured goods and the real exchange rate. Now, combining a log-differentiated version of Eq. (13) with Eq. (14), we can derive Eq. (15), which demonstrates a decline in the home country's manufactured exports in response to rising global commodity prices, with the size of damage positively depending on the degree of real appreciation:

$$\frac{\partial \hat{C}_M^*}{\partial \hat{p}_R^*} = \varepsilon^d \eta \left(\frac{\partial \hat{Q}}{\partial \hat{p}_R^*}\right) \leq 0, \tag{15}$$

where $\partial \hat{Q} / \partial \hat{p}_R^* > 0$ by Eq. (11). ■

An intuitive interpretation of Eq. (15) is that a surge in commodity prices is expected to increase the domestic input costs (i.e., wage rates) of producing manufactured goods and squeeze manufacturers' profits, thereby reducing their incentives for production. The lower supply is then followed by a rise in the price of manufactured exports, adversely affecting the foreign demand.

Proposition 3. *Capital controls restrict the magnitude of the real exchange rate response to a commodity price shock.*

Proof. Deviating from the benchmark model assumption, let us now consider an extreme case of capital market autarky to study the effect of capital controls. With no cross-border capital flows, the return to capital r is endogenously determined in the domestic capital market. Resolving the model with only domestically mobile capital and labor, we find the following real exchange rate response to a commodity price shock:

$$\frac{\partial \hat{Q}}{\partial \hat{p}_R^*} = \varphi\theta > 0 \tag{16}$$

$$\text{with } \varphi = 1/\left(\mu_{L,R} - \left(\mu_{K,R}/\mu_{K,M}\right)\mu_{L,M}\right),$$

where $\mu_{K,i}$ is the capital income share $\left(0 < \mu_{K,i} < 1\right)$, defined as $\mu_{K,i} \equiv rK_i/p_iY_i$ in sector i. By comparing Eqs. (11) and (16), we observe that the real exchange rate reaction is smaller in the presence of capital controls because $\varphi < \tau$.[10] ■

This result occurs because a given rise in commodity prices boosts the rental rate for capital as well as the wage rate when cross-border capital movement is restricted, making the resulting increase in the wage rate smaller than would be the case with free international capital mobility (see Eqs. (E.7) and (E.13) in Appendix E). As a result, the price of nontraded goods will increase less under the capital control, mitigating the appreciation pressures of the real exchange rate.

[10] Note that $\mu_{K,R} < \mu_{K,M}$ due to the assumption in Footnote 7.

2.5. Testable hypotheses

Combining Propositions 1 and 3 above, our first testable hypothesis is:

Hypothesis 1. *Capital controls lessen the transmission of commodity price shocks to the real exchange rate*.

Moreover, combining Propositions 2 and 3 above, the second testable hypothesis is:

Hypothesis 2. *Capital controls lower the propensity to crowd out manufactured exports arising from a commodity price boom*.

In the next section, we build empirical models to test the above two hypotheses using a panel dataset.

3. Data and empirical model specification

Our sample covers 37 non-oil commodity-exporting countries for the period of 1980−2017. See Appendix A for a full list of sample countries. Major energy exporters, especially oil exporters, are not part of our sample because of their highly volatile export prices and various strategic pricing behaviors (e.g., possible collusion among OPEC countries), which can complicate their economies' transmission mechanisms between resource export prices and real exchange rates. In fact, almost all of the large oil-exporting countries peg their currencies to the dollar and do not allow nominal exchange rate adjustments to an external shock.

For our purpose, we keep commodity-dependent countries with a non-negligible share of manufactured exports, so in the vast majority of our sample countries, at least 5% of their total exports are manufactured products.

In the rest of the section, we briefly explain the definition and source of the variables used in our empirical analysis and then present the baseline regression models.

3.1. Key variables

3.1.1. Real exchange rate

We use the CPI-based real effective exchange rate, 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. It is measured such that the higher index indicates the real appreciation of the domestic currency. The monthly and annual real effective exchange rate series are taken from the Bruegel database released by Darvas (2012).

3.1.2. Real commodity price

The real commodity price index is defined as the world (nominal) price index of a country's commodity exports relative to the world price index of manufactured goods exports. Following Cashin et al. (2004) and Chen and Lee (2018), we construct a country-specific real commodity price index using 58 commodities as follows:[11]

[11] For a complete list of commodities, see Appendix A.

$$RCP_{it} = \left[\sum_{j=1}^{J} w_{ij}\left(\ln p_{jt}\right)\right]/MUV_t \tag{17}$$

$$\text{with } w_{ij} = \left(1/T\sum_{t=1}^{T} ex_{ij,t}\right)/\left(1/T\sum_{t=1}^{T} EX_{it}\right)$$

where p_{jt} is the global price of commodity j at time t, MUV_t is the unit value index of manufactured exports for 20 industrial economies, $ex_{ij,t}$ is country i's export volume (in U.S. dollars) of commodity j, and EX_{it} is the volume of the total commodity exports of country i. We keep weight w_{ij} constant over time to eliminate the quantity effect from the price index calculation.[12] Whenever necessary, we take the average of monthly commodity price indices in each year to convert them to an annual frequency.

The monthly world commodity price series are extracted from the International Monetary Fund (IMF) and World Bank's Pink Sheet data, the unit value index of manufactured exports from the IMF's International Financial Statistics, and the annual commodity trade data from the UN COMTRADE database.

3.1.3. Capital controls

For the baseline regression analysis, we build a capital control variable based on an annual de facto international financial integration taken from the updated External Wealth of Nations Mark II database available in Lane and Milesi-Ferretti (2017). Among the integration indicators proposed by Lane and Milesi-Ferretti (2003), we adopt a measure of cross-border equity holdings that is defined as follows:

$$GEQ_{it} = \left(EQ_{it}^{A} + FDI_{it}^{A} + EQ_{it}^{L} + FDI_{it}^{L}\right)/GDP_{it} \tag{18}$$

[12] More specifically, we use the period-average values of export volume of each commodity over the period of 1986−2010.

where EQ_{it} and FDI_{it} are respectively country i's stocks of portfolio equity and foreign direct investment at time *t*, with the superscript A indicating assets, and the superscript L, liabilities. A higher value of GEQ in Eq. (18) represents a more open capital account.

We limit our attention to the equity-based measure to be broadly consistent with the model environment in our theoretical framework, excluding debt instruments and foreign exchange reserves. In fact, as noted by Kose et al. (2009), debt flows tend to be highly volatile and can magnify the negative impact of adverse shocks in developing economies.[13] In order to create a capital control indicator, we take the inverse of GEQ so that a higher value of the indicator (= $1/GEQ$) corresponds to stricter restrictions on capital flows.[14]

The considerable time variation for de facto capital controls at the country level makes them preferable to de jure measures, as it helps identify the intended effect of capital market regulations in our panel fixed-effect regressions. The de facto indicators also allow us to distinguish between controls on capital inflows and controls on capital outflows during the sample period.

3.1.4. Manufactured exports

We use manufactured exports as a share of GDP. The annual data are taken from the World Bank's World Development Indicators (WDI).

[13] Kose et al. (2009) also acknowledge that de facto financial openness measures tend to better capture the extent of a country's integration into global financial markets than de jure ones because the latter cannot capture the degree of enforcement and effectiveness of capital controls.

[14] To mitigate the influence of outliers, we drop the top 1% (inclusive) of observations for capital controls before conducting a regression analysis.

3.2. Other variables

Other control variables in our empirical analysis include *government spending* (the log of the ratio of government consumption to GDP), *trade openness* (the log of the sum of exports and imports relative to GDP), and *investment* (the log of the ratio of gross capital formation to GDP). We obtain the information for these variables from the World Bank's WDI.

In addition, since sectoral output and employment data are not available for the bulk of our sample countries, we follow Lane and Milesi-Ferretti (2004) and define *relative GDP per capita* as the trade-weighted sum of the log of the home country's GDP per capita relative to its trading partners'. It is included to capture relative output levels and control for a Balassa–Samuelson effect in the real exchange rate regressions. Bilateral trade data are collected from the IMF's Direction of Trade Statistics and GDP per capita in constant 2010 U.S. dollars from the World Bank's WDI.

Lastly, in order to control for the effect of foreign demand in the manufactured export regressions, we create *foreign income* as the trade-weighted sum of the log of trading partners' GDP per capita. Summary statistics for all variables are presented in Appendix Table B.1.

3.3. Baseline regression model specifications

As a preliminary procedure, we apply the standard panel time-series tests to our dataset and find the presence of non-stationarity for all annual variables including the real exchange rate and real commodity price indices. We also find evidence of cointegration among the annual variables at the conventional significance level (results available in Appendix Tables C.1 and

C.2). Accordingly, we employ a panel version of the dynamic ordinary least squares (DOLS) estimator to efficiently estimate the long-run cointegrating relationship, which uses a parametric correction for endogeneity by including the leads and lags of the first difference of each regressor.[15]

Kao and Chiang (2000) provide evidence that DOLS is superior to the fully modified ordinary least squares (FMOLS) estimator, another widely used methodology, 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 would have to rely on a substantially reduced sample size.

For country *i* and year *t*, the first baseline regression model takes the following panel DOLS(1,1) specification:

$$\begin{aligned} RER_{it} &= \alpha_1 RCP_{it} + \alpha_2 (RCP_{it} \times KC_{it}) + \alpha_3 KC_{it} + \boldsymbol{X_{it}\gamma} \\ &+ \textstyle\sum_{j=-1}^{1} \Delta \boldsymbol{Z_{i,t+j}\delta_j} + \phi_i + \phi_t + \varepsilon_{it} \end{aligned} \tag{19}$$

where RER_{it} is the log of the real effective exchange rate; RCP_{it} is the log of the real commodity price index; KC_{it} is a measure of capital control; $\boldsymbol{X_{it}}$ is a vector of additional fundamental determinants, including *government spending*, *relative GDP per capita*, and *trade openness*; $\boldsymbol{Z_{it}}$ is a vector of all continuous explanatory variables; ϕ_i is a country fixed effect; ϕ_t is a time fixed effect; ε_{it} is a residual; and Δ is the first-difference operator. By controlling for country and time fixed effects, the problem of omitted variables bias or misspecification is diminished. To account

[15] As noted by Lane and Milesi-Ferretti (2004), "the superconsistency property of cointegrated equations means that any possible endogeneity running from the real exchange rate to the regressors does not affect the estimated long-run coefficients."

for potential cross-sectional correlation as well as autocorrelation and heteroscedasticity, we use Driscoll and Kraay's (1998) standard errors for statistical inferences.

Our Hypothesis 1 tests whether $\alpha_1 > 0$ and $\alpha_2 < 0$ in Eq. (19) so that the positive impact of *RCP* shock on *RER* (or the *RCP* elasticity of *RER*) may be reduced through restrictions on cross-border capital movements. Regarding other control variables, government consumption is typically spent on nontraded goods, and we expect a positive coefficient for *government spending*. Due to the Balassa–Samuelson effect, *relative GDP per capita* is expected to enter the *RER* regression with a positive sign. *Trade openness* tends to increase the share of tradable goods in domestic consumption, so we expect it to have a negative effect on *RER*.

The second baseline regression model takes the following panel fixed-effect estimator:

$$MX_{it} = \beta_1 RCP_{it} + \beta_2 (RCP_{it} \times KC_{it}) + \beta_3 KC_{it} + \boldsymbol{Y_{it}} \boldsymbol{\gamma} + \phi_i + \phi_t + e_{it} \tag{20}$$

where MX_{it} is the log of the ratio of manufactured exports to GDP in country *i* at time *t*, and $\boldsymbol{Y_{it}}$ is a vector of other potential determinants of country *i*'s exports of manufacturing, including *trade openness*, *investment*, and *foreign income*.

In order to focus on the long-run effects of commodity price movements on manufactured exports, we smooth out the business cycle fluctuations by transforming the annual frequency data into five-year averages, as is standard in the growth literature (e.g., Rodrik, 2008; Aghion et al., 2009).

Our Hypothesis 2 tests whether $\beta_1 < 0$ and $\beta_2 > 0$ in Eq. (20) so that the negative impact of *RCP* shock on *MX* may be moderated through restrictions on international capital movements. Regarding the other regressors, a greater value of *investment* is likely to promote *MX* owing to an increase in available physical capital, which may be required for manufacturing production. The higher level of *trade openness* is usually associated with lower trade barriers in tariffs and quotas, likely boosting a country's foreign trade, including *MX*. The demand for domestically produced manufactured goods would increase with trading partners' purchasing power, so *foreign income* is expected to show a positive sign.

4. Empirical results

4.1. Main results

Columns (1)−(3) of Table 1 present the estimation results based on our first baseline regression model in Eq. (19). The main parameters of our interest are on the coefficients of the commodity price index *RCP* and its interaction with capital controls $RCP \times KC$.

[Insert Table 1 here]

Column (1) displays a significantly positive coefficient for *RCP*, which demonstrates its long-run cointegrating relationship with *RER* in our sample countries. This result reinforces the previous empirical evidence for the commodity currency phenomenon documented in Chen and Rogoff (2003), Cashin et al. (2004), Coudert et al. (2011), Ricci et al. (2013), Bodart et al. (2012, 2015), and Chen and Lee (2018).

Column (2) extends the specification with additional fundamental determinants of *RER*, including *government spending*, *relative GDP per capita*, and *trade openness*. We confirm a positive long-run relationship between *RCP* and *RER*, with expected signs for the other control variables. Indeed, inclusion of the other *RER* determinants strengthens the magnitude of *RCP* elasticity and its statistical significance.

In column (3), we further extend the model with *KC* and its interaction with *RCP*.[16] Significantly positive *RCP* and negative *RCP* × *KC* coefficient estimates indicate that while an increase in commodity prices induces real appreciation, a more stringent capital control (a higher value of *KC*) appears to reduce the size of appreciation, in support of our Hypothesis 1. In particular, a 1% rise in *RCP* would lead to long-run real appreciation of 0.56% when *KC* is at its sample average and appreciation of 0.37% when there is a one–standard deviation increase in *KC* above its mean value.[17]

Turning to the *MX* regressions, we first show in column (4) a statistically significant and negative response of *MX* to *RER* appreciation, consistent with the conventional theory. The negative coefficient estimate of *RER* indicates that a 1% increase in *RER* tends to lower *MX* by 0.72% in our sample countries.

We now introduce *RCP* as a determinant of *MX* while controlling for other relevant variables. As shown in column (5), a significantly negative *RCP* coefficient provides empirical evidence for the Dutch disease, the coexistence of a commodity boom and manufacturing shrinkage, in commodity-exporting developing countries.[18] Other control variables such as *trade*

[16] Note that the source data for *KC*, the updated External Wealth of Nations Mark II database (Lane and Milesi-Ferretti, 2017), is available up to 2015, so the specification that includes *KC* has a smaller sample size.

[17] The net effects of a 1% increase in *RCP* are calculated by $\alpha_1 + (\alpha_2 \times \text{mean}_{KC})$ and $\alpha_1 + \left(\alpha_2 \times (\text{mean}_{KC} + \sigma_{KC})\right)$, respectively.

[18] We have also considered a specification that includes both *RER* and *RCP* at the same time to test whether the former drives out the effect of the latter in the *MX* regression. The estimation results, available upon request, show that both variables keep their expected negative signs, but only *RER* remains strongly significant. This result verifies the role of *RER* as an intermediate channel through which an *RCP* boom may hurt *MX* in developing countries.

openness, *investment*, and *foreign income* have the expected positive signs, although *foreign income* is not significant at standard confidence levels.

Finally in column (6), we have a full specification, as in our second baseline regression model in Eq. (20). A negative *RCP* coefficient and a positive coefficient for the interaction term between *RCP* and *KC* lend support to our Hypothesis 2. Specifically, a 1% rise in *RCP* would decrease *MX* by 1.89% when *KC* is at its sample average and by 1.73% when *KC* is at one standard deviation above its mean value. In other words, capital flow regulations are expected to slow down a manufacturing downturn in developing countries by resisting the appreciation pressures associated with a commodity price boom.

The result in column (6) also shows that *KC* itself has a negative effect on *MX*, although it is only marginally significant. Some plausible explanations for this result are as follows: higher barriers on capital mobility can contract manufacturing production through a limited supply of inputs in the foreign capital-dependent production process, or through foregone opportunities to benefit from positive spillovers generated by FDI in the commodity sector. While the net effect of tighter *KC* on *MX* is positive in our sample, the negative standalone effect of *KC* suggests that a careful cost–benefit analysis across industries may precede the imposition of *KC* to exploit foreign capital more effectively.[19]

In addition to individual coefficient estimates and their standard errors, Table 1 also reports *p*-values for *F*-statistics to test the null hypothesis that *RCP* has no effect on *RER* and *MX* in the interaction variable regressions. As seen in Eqs. (19) and (20), this null hypothesis requires a joint significance test for *RCP* and its interaction with *KC*. The consistently low *p*-values reported in columns (3) and (6) validate our baseline empirical specifications. Likewise, the

[19] Using the result in column (6) of Table 1, the net effect of *KC* on *MX* can be evaluated by $\{\exp[(\beta_2 \times \text{mean}_{RCP} \times \sigma_{KC}) + (\beta_3 \times \sigma_{KC})] - 1\} \times 100$.

relatively low *p*-values for a joint significance test for *KC* and its interaction with *RCP* provide further support for the validity of our specifications.

4.2. Alternative capital control indicators

In this subsection, we test whether our main results are sensitive to alternative measures of capital controls. As a first exercise, we use Chinn and Ito's (2006) index, which is one of the most widely used de jure measures of capital account openness. It is built upon the information about legal or regulatory barriers to international financial transactions reported in the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions. As higher values of the index represent more open capital markets, we define a capital control dummy variable that takes a value of unity at time *t* if the Chinn–Ito index for a country is below the 20th percentile in our sample and zero otherwise.[20]

In a second exercise, we employ the KOF hybrid financial globalization index, available at the KOF Swiss Economic Institute (Gygli et al., 2019), which combines de facto and de jure indices with equal weights.[21] The de facto index is based on work of Lane and Milesi-Ferretti (2007, 2017) and takes a quantity-based measure of stocks of foreign assets and liabilities. More specifically, it consists of 27.6% international debt, 27.1% international income payments, 26.7% FDI, 16.5% portfolio investment, and 2.1% international reserves. On the other hand, the de jure index is based on the indicator developed by Chinn and Ito (2006) and the investment restrictions published in the World Economic Forum Global Competitiveness Report. It is composed of 38.5% capital account openness, 33.3% investment restrictions, and 28.2% international investment

[20] We have also considered the 10th and 30th percentiles as alternative thresholds and found very similar results.
[21] The original KOF globalization index was introduced by Dreher (2006) and later updated by Dreher et al. (2008).

agreements. Since a higher value of the index represents that an economy is more financially globalized, we use the inverse of the KOF hybrid index as a measure of capital controls.

Table 2 reports the estimation results when we construct *KC* based on the Chinn–Ito index in columns (1) and (3) and the KOF hybrid index in columns (2) and (4). Indeed, we find that the interaction effect between *RCP* and *KC* retains the expected signs in all cases, though it is not always statistically significant (the *p*-value for the interaction term is 0.53 in column (3)). One of the reasons for the difficulty of identifying the interaction effect in column (3) is the relatively little time variation in the Chinn–Ito index at the country level.

[Insert Table 2 here]

4.3. Robustness test controlling for exchange rate regimes and financial crises

In order to test the robustness of our main results, we introduce two more factors into the baseline regression models. The goal is to see if the variable of our main interest, the interaction of *RCP* and *KC*, continues to play an important role when controlling for other variables that might affect the transmission of *RCP* changes to *RER* and *MX*.

The first variable we add is a country's choice of a fixed vs. a flexible exchange rate regime. To do so, we follow Ilzetzki et al. (2019) and define a "*flexible regime*" dummy variable using their fine classification code. This dummy takes a value of one in a given year if the code for a country is between 5 and 14, or zero if the code is below 5. In the case of five-year average data, we first take the average of classification codes and then generate a binary regime variable following the same rule. By construction, the reference category (i.e., *flexible regime* = 0) is a de facto peg or preannounced horizontal band with margins of no larger than ±2%.[22]

From a theoretical point of view, even if the nominal exchange rate remains fixed in pegged countries, a more stable real exchange rate in the long run will not be guaranteed because a priori, we do not know how much domestic prices will react to spikes in commodity prices relative to the reaction in nonpegged countries. For this reason, the impact of exchange rate regimes is more of an empirical issue that deserves further investigation.

Columns (1) and (3) of Table 3 show the regression results when *flexible regime* and its interaction with *RCP* are included as additional controls. First of all, we continue to see the expected signs, with strong significance for the *RCP* and *KC* interaction variable, although the inclusion of multiple interaction variables that may be highly correlated lessens the statistical significance of the estimates for some of the regressors.

[Insert Table 3 here]

Moreover, in column (1), we find a significantly negative sign for *RCP*'s interaction with *flexible regime*. This result reflects that a flexible nominal exchange rate provides a more effective *RER*-stabilizing role in the long run for a country facing a commodity price boom, in accordance with the findings of Bodart et al. (2015). Nevertheless, the interaction between *RCP* and *flexible regime* does not necessarily help shield manufactured exports, as its coefficient estimate in column (3) has a negative sign although it is not statistically significant.

[22] We exclude episodes of "Dual market in which parallel market data is missing" (fine classification code = 15) from the sample for regression analysis.

The second factor we introduce is a major financial crisis that developing countries in our sample have undergone during the sample period. We create a "*crisis*" variable that reflects country-level banking crises as well as the 2008−09 global financial crisis and define it as the sum of crisis years divided by the number of years in the corresponding period. Hence, for the annual data, *crisis* is a dummy variable to capture a crisis year. It intends to capture severe financial market instability that has the potential to cause large changes for our dependent variables. The information for the banking crisis years comes from the World Bank's Global Financial Development database.

Columns (2) and (4) of Table 3 report the results when controlling for the interaction between *RCP* and *crisis*. While we find no significant effects of the financial crisis on the transmission of *RCP* changes into *RER* and *MX*, the *RCP* and *KC* interaction effects stay significant with the expected signs, confirming the robustness of our main results.

4.4. Quantile regression evidence for nonlinearity

The main focus of our analysis is on the Dutch disease resulting from a commodity price boom, so we looked into two relationships in Tables 1−3: one between *RCP* and *RER*, and the other between *RCP* and *MX*, with *KC* playing a dampening role in both relationships. By the model's design, the operative channel through which a country suffers from the Dutch disease is the extent of its real appreciation.

In this subsection, we test the possible nonlinearity between commodity currency responses and their impact on manufactured exports using the following quantile regression model:

$$MX_{it} = \sum_{k=1} \delta_k (RCP_{it} \times \mathbf{1}[\underline{\mathrm{p}} \leq E_{it} < \overline{\mathrm{p}}]) + \boldsymbol{Y_{it}\gamma} + \phi_i + \phi_t + u_{it} \tag{21}$$

where the indicator function $\mathbf{1}[\cdot]$ takes the value of one when the commodity price elasticity E_{it} for country i at time t falls within a specified percentile range. The country-specific commodity price elasticity is estimated by DOLS(1,1) using monthly *RER* and *RCP* for five-year periods.

Our conjecture is that the more sensitive the *RER* response is to *RCP* changes (i.e., the larger the elasticity), the greater the crowding-out effect of the commodity price boom on manufactured exports due to a larger loss of competitiveness.

Table 4 displays the estimation results based on the model in Eq. (21) with and without other control variables in columns (1) and (2). Consistent with the conjecture above, we find robust empirical evidence for more drastic reductions in manufactured exports as the commodity price elasticity grows. For example, the results in column (2) suggest that a 1% increase in *RCP* is expected to lower *MX* by 2.36% on average when the elasticity falls below the 33rd percentile in its distribution, by 2.69% when it is between the 33rd and 66th percentiles, and by 3.14% when it exceeds the 66th percentile.

[Insert Table 4 here]

A more general pattern is illustrated in Fig. 2. In panel (a), we plot the marginal effects of commodity prices on manufactured exports in finer elasticity quantiles. Panel (b) plots the fitted values of *MX* in various elasticity quantiles when *RCP* takes its sample average. The concave-downward slope in both plots indicates that when *RER* is more sensitive to *RCP* movements, there is a more severe crowding-out effect on *MX* given a commodity price shock.[23] This observation, in combination with the main results in Table 1, suggests the countercyclical use of capital controls in countries whose currency values strongly co-move with their commodity export prices in order to protect non-commodity tradable sectors.

[Insert Fig. 2 here]

4.5. Real exchange rate misalignments and manufactured exports

Although our interpretations have focused on the case of real appreciations, the results presented thus far do not reveal a possible asymmetry in *MX* responses following changes in *RER*. We thus investigate the cases for under- and overvaluations of *RER* relative to its equilibrium levels and their possibly different impacts on *MX*. Three versions of *RER* misalignments are considered here.

Our first approach is to follow Rodrik (2008) and define a misalignment as a difference between the actual *RER* and the rate adjusted for the Balassa–Samuelson effect based on a pooled regression. Specifically, we regress *RER* on *relative GDP per capita* and a time fixed effect. We then subtract the fitted value from the actual *RER* to arrive at the overvaluation if the difference is greater than zero and the undervaluation if it is smaller than zero.

Our second approach is to calculate the misalignment series as the departures of the actual *RER* from a Hodrick–Prescott (H-P) filtered series that represents an estimated

[23] Fig. D.1 in the Appendix displays a similar pattern when the estimations are performed using a model that also controls for the other macroeconomic determinants of *MX*.

equilibrium *RER*. As in Goldfajn and Valdés (1999), the H-P filter-based misalignment (MIS) for each country is computed as follows:

$$MIS_t = 100 + 100 \times (RER_t - \overline{RER}_t)/\overline{RER}_t \tag{22}$$

where $\overline{RER}_t$ is the H-P filtered series. From Eq. (22), we can see that the estimated misalignment series captures the cyclical component of the *RER* movements and takes a value greater than 100 for overvaluation and less than 100 for undervaluation.

Our third approach is to find the predicted *RER* for each country based on the cointegrating relationship between *RER* and a set of nonstationary fundamentals such as *RCP*, *government spending*, *relative GDP per capita*, and *trade openness*. We then use Eq. (22) with $\overline{RER}_t$ being the fitted *RER* series to calculate fundamental-based misalignment series. Note that, like Goldfajn and Valdés (1999), we use H-P filtered fundamentals to calculate the fitted *RER*.

To test whether *RER* misalignments would crowd out manufactured exports, Table 5 sets out the estimation results with overvaluation in the upper panel and undervaluation in the lower panel.

[Insert Table 5 here]

The upper panel of Table 5 reports significant and robust evidence for a negative impact of overvaluation on manufactured exports, with a misalignment calculation accounting for the Balassa−Samuelson effect in column (1), a H-P filtered equilibrium in column (2), and cointegrated fundamentals in column (3). These results are consistent with those of Prasad et al. (2007), who also emphasize a negative association between real overvaluation and the growth of exportable manufacturing sectors.

By contrast, the results in the lower panel show no consistent patterns of statistical significance or coefficient sign, suggesting that *RER* undervaluation may not have a definite effect on manufactured exports. Overall, a central lesson we learn from the results in Table 5 is that excessive real appreciation is key to deterring export promotion in the manufacturing sector.

4.6. Evidence from different types of capital controls

The *KC* variable used in the analysis in Tables 1 and 3 is an index that uses the information for cross-holdings of portfolio equity and direct investment combined. As an aggregate measure, it does not distinguish capital inflows from outflows or portfolio equity from FDI flows. To identify the primary driving forces behind the dampening role of capital controls, we disaggregate the *KC* variable into FDI vs. portfolio equity and outward vs. inward for each asset category.

We first generate the following financial integration indicators using the External Wealth of Nations dataset (Lane and Milesi-Ferretti, 2017): FDI overall, FDI inward, FDI outward, (portfolio) equity overall, equity inward, equity outward, GEQ inward, and GEQ outward.[24] Inward (outward) indicators are defined as the ratio of the liabilities (assets) of the corresponding capital categories to GDP, and overall indicators as the sum of inward and outward indicators. We then follow the procedure in Section 3.1.3 and create a proxy for capital controls by taking the inverse of the financial integration indicators for either direction for each category.

Table 6 summarizes the results when we redefine the *KC* variable at the disaggregate level with *RER* as the dependent variable in columns (1)−(5) and *MX* as the dependent variable in columns (6)−(10). As you may notice, we do not report the results with outward indicators, as

[24] GEQ overall is what we have used as the baseline measure of *KC*.

all estimation results that involve them are less statistically and economically significant (results available in Appendix Table D.1). This is in line with our findings in Table 5 in that *RER* overvaluation is more of a concern than undervaluation, and overvaluation is more related to inward, rather than outward, capital movements.

[Insert Table 6 here]

Reviewing the results for *RER* regressions with FDI regulations between columns (1) and (2), we find that the magnitude and significance levels of coefficient estimates are very similar. The same is true for the results for *MX* regressions between columns (6) and (7).

When looking at the results in columns (3), (4), (8), and (9), we find little evidence for a strong effect of portfolio equity flow regulations; even if the coefficient estimates of the interaction term are statistically significant, their magnitude is too small to have any meaningful economic impacts.[25] This is not surprising because stock markets in our sample countries represent a relatively small fraction of the domestic economy.

Furthermore, we see that the results in columns (5) and (10) are very close to those in columns (3) and (6) of Table 1, confirming the patterns we observed between FDI overall and inward regulations from Table 6. The main message emerging from these results is that restrictions on inward FDI are mostly responsible for reducing *RCP*'s transmissions to *RER* and *MX* in the long run in commodity-dependent developing countries.

[25] Note also that due to the missing observations for portfolio equity in some of our sample countries, regressions in columns (3) and (4) rely on 35 countries, and those in (8) and (9) on only 34 countries.

5. Conclusion

Slow economic growth in developing countries that rely heavily on raw commodity products has been a long-standing topic in economics. Indeed, the empirical literature on the Dutch disease extensively documents that while commodity windfall gains have positive short-run impacts on economic growth, their long-term effects tend to be negative. Unsurprisingly, even if a country has a comparative advantage in producing primary commodities, it may have an incentive to expand the manufacturing sector, which can provide momentum for long-run growth due to learning-by-doing and knowledge spillovers (van Wijnbergen, 1984; Krugman, 1987; Matsuyama, 1992; Sachs and Warner, 1995; Gylfason et al., 1999; Torvik, 2001).

How then can commodity-abundant countries promote their economic diversification? We address this question with a particular focus on the merits of capital controls in stabilizing real exchange rates and alleviating the intensity of the Dutch disease in response to a sharp increase in commodity prices.

Consistent with the theory-based hypotheses, we find significant evidence that there is a strong positive association between real exchange rates and commodity export prices in the long run, with the extent of this relation weaker when the cross-border capital flows, particularly of inward FDI, are more strictly regulated. Capital controls in turn seem to attenuate the propensity to crowd out manufactured exports by reducing real appreciation pressures following a surge in commodity prices.

Our results highlight the importance of countercyclical capital controls to lessen the adverse effects of terms-of-trade movements on the exchange rate and trade, thereby accelerating export diversification and industrialization in resource-rich developing countries.

We acknowledge that exchange rate stabilization through capital account managements is not the only industrialization policy available in commodity-dependent countries. Policies facilitating investments in infrastructure, education, and R&D can also encourage production of the manufacturing sectors and complement capital controls to further enhance growth potential.

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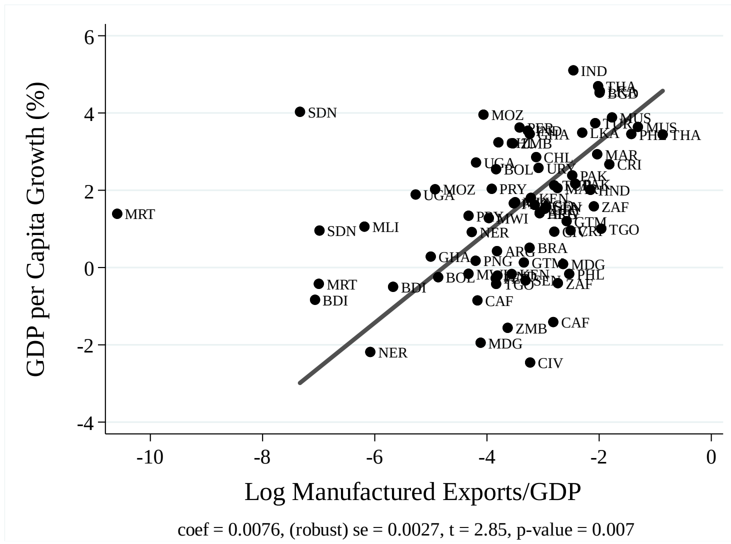
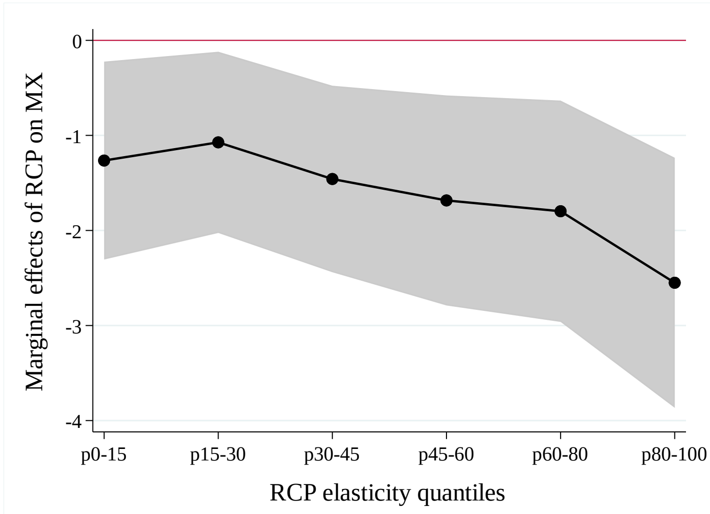
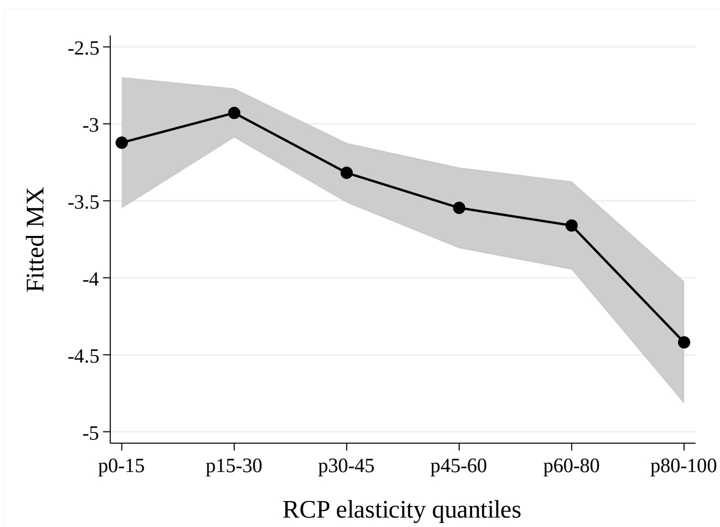


Fig. 1. Manufactured exports and GDP per capita growth, 1980–2017.

Notes: To obtain the fitted values, the growth rate of per capita GDP is regressed on manufactured exports, primary product (natural resource) exports, government spending, investment, trade openness, secondary schooling, population growth, country and time fixed effects (all in logs except for the last three variables). Data source: World Bank's WDI.



(a) Marginal effects of commodity prices on manufactured exports



(b) Predicted manufactured exports

Fig. 2. Marginal effects of commodity prices and predicted manufactured exports.

Notes: *RCP* = (log) real commodity price; *MX* = (log) manufactured exports/GDP. In panel (a), we plot the marginal effects of commodity prices, $\partial MX/\partial RCP$, based on the model $MX_{it} = \sum_{k=1} \delta_k (RCP_{it} \times \mathbf{1}[\cdot]) + \phi_i + \phi_t + u_{it}$, where the indicator function $\mathbf{1}[\cdot]$ takes the value of one when the commodity price elasticity of real exchange rate for country *i* at time *t* falls within a specified percentile range. Using the same model, panel (b) illustrates the predicted or fitted values of *MX* when *RCP* takes its sample average. The gray bands in both graphs represent 90% confidence intervals.

Table 1
The impact of commodity prices and capital controls: main results.

Estimation method	DOLS(1,1)			Panel FE		
Dependent variable	Real exchange rate			Manufactured exports		
	(1)	(2)	(3)	(4)	(5)	(6)
RCP	0.524**	0.828***	0.697**		-2.224**	-2.004*
	(0.194)	(0.256)	(0.271)		(0.793)	(0.884)
RCP × *KC*			-0.010***			0.009**
			(0.003)			(0.003)
KC			0.008**			-0.009*
			(0.003)			(0.004)
Government spending		0.177**	0.223***			
		(0.079)	(0.061)			
Relative GDP per capita		1.778***	2.112***			
		(0.338)	(0.374)			
Trade openness		-0.447***	-0.465***		1.146***	1.121***
		(0.134)	(0.128)		(0.108)	(0.111)
Investment					0.313**	0.345**
					(0.104)	(0.118)
Foreign income					0.057	0.048
					(0.071)	(0.073)
RER				-0.719***		
				(0.138)		
Observations	1,154	1,060	983	255	268	268
R^2 (within)	0.166	0.365	0.439	0.293	0.426	0.431
p-values for joint significance						
RCP and *RCP* × *KC*			< 0.01			0.015
KC and *RCP* × *KC*			< 0.01			0.065

Notes: *RCP* = real commodity price; *KC* = capital control; *RER* = real exchange rate. All variables in the table are measured in logarithms except for *KC*. DOLS(1,1) procedure includes contemporaneous, 1 lead, and 1 lag of changes of all continuous regressors, but they are suppressed to save space. All specifications include both country and time fixed effects. Driscoll-Kraay standard errors are reported in parentheses. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. The estimations are performed based on annual observations in columns (1)−(3) and non-overlapping five-year averages in columns (4)−(6).

Table 2
The impact of commodity prices and capital controls: using alternative capital control measures.

Estimation method	DOLS(1,1)		Panel FE	
Dependent variable	Real exchange rate		Manufactured exports	
Source of capital controls	Chinn-Ito	Hybrid	Chinn-Ito	Hybrid
	(1)	(2)	(3)	(4)
RCP	0.780***	0.765**	-2.158**	-2.250**
	(0.267)	(0.287)	(0.859)	(0.782)
RCP × *KC*	-0.160***	-0.111+	0.062	0.343**
	(0.036)	(0.075)	(0.093)	(0.129)
KC		-0.049		-0.071
		(0.077)		(0.096)
Government spending	0.121+	0.167**		
	(0.075)	(0.073)		
Relative GDP per capita	1.413***	1.436***		
	(0.362)	(0.391)		
Trade openness	-0.443***	-0.550***	1.160***	1.328***
	(0.136)	(0.151)	(0.099)	(0.173)
Investment			0.301**	0.398**
			(0.097)	(0.136)
Foreign income			0.053	0.049
			(0.071)	(0.074)
Observations	1,060	1,046	268	266
R^2 (within)	0.401	0.387	0.427	0.444
p-values for joint significance				
RCP and *RCP* × *KC*	< 0.01	0.017	0.015	< 0.01
KC and *RCP* × *KC*		0.029		0.011

Notes: *RCP* = real commodity price; *KC* = capital control. All variables in the table are measured in logarithms except for *KC*. DOLS(1,1) procedure includes contemporaneous, 1 lead, and 1 lag of changes of all continuous regressors, but they are suppressed to save space. All specifications include both country and time fixed effects. Driscoll-Kraay standard errors are reported in parentheses. ***, **, and + indicate statistical significance at the 1%, 5%, and 15% levels, respectively. The estimations are performed based on annual observations in columns (1) and (2) and non-overlapping five-year averages in columns (3) and (4).

Table 3
Robustness checks controlling for exchange rate regimes and financial crises.

Estimation method	DOLS(1,1)		Panel FE	
Dependent variable	Real exchange rate		Manufactured exports	
	(1)	(2)	(3)	(4)
RCP	0.800***	1.057***	-1.483	-2.118***
	(0.279)	(0.222)	(1.148)	(0.554)
RCP × *KC*	-0.012***	-0.005*	0.014**	0.008**
	(0.003)	(0.003)	(0.005)	(0.003)
RCP × *Flexible regime*	-0.183**		-0.472	
	(0.077)		(0.787)	
RCP × *Crisis*		0.104		-0.607
		(0.113)		(0.481)
KC	0.010***	0.003	-0.013**	-0.013**
	(0.003)	(0.003)	(0.005)	(0.005)
Flexible regime	0.086		0.473	
	(0.083)		(0.736)	
Crisis		-0.076		0.284
		(0.106)		(0.536)
Government spending	0.223***	0.213***		
	(0.056)	(0.060)		
Relative GDP per capita	2.168***	1.546***		
	(0.397)	(0.178)		
Trade openness	-0.463***	-0.438***	1.149***	1.286***
	(0.132)	(0.121)	(0.119)	(0.116)
Investment			0.331**	0.173
			(0.129)	(0.131)
Foreign income			0.064	0.262***
			(0.086)	(0.051)
Observations	982	983	266	268
R^2 (within)	0.448	0.384	0.433	0.419
p-values for joint significance				
RCP and *RCP* × *KC*	< 0.01	< 0.01	0.034	< 0.01
KC and *RCP* × *KC*	< 0.01	< 0.01	0.052	0.094

Notes: *RCP* = real commodity price; *KC* = capital control. All variables in the table are measured in logarithms except for *KC*, *Flexible regime*, and *Crisis*. DOLS(1,1) procedure includes contemporaneous, 1 lead, and 1 lag of changes of all continuous regressors, but they are suppressed to save space. All specifications include country fixed effects. Driscoll-Kraay standard errors are reported in parentheses. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. The estimations are performed based on annual observations in columns (1) and (2) and non-overlapping five-year averages in columns (3) and (4).

Table 4
The impact of the larger commodity price elasticity.

Dependent variable	Manufactured exports	
	(1)	(2)
RCP × **1**[p0 ≤ *E* < p33]	-1.451+	-2.355**
	(0.786)	(0.970)
RCP × **1**[p33 ≤ *E* < p66]	-1.960**	-2.694**
	(0.792)	(0.973)
RCP × **1**[p66 ≤ *E*]	-2.464**	-3.144**
	(0.965)	(1.155)
Trade openness		1.122***
		(0.141)
Investment		0.280**
		(0.100)
Foreign income		0.092
		(0.088)
Observations	255	248
R^2 (within)	0.283	0.385

Notes: *RCP* = real commodity price. The indicator function **1**[·] takes the value of one when the commodity price elasticity of real exchange rate *E* falls within a specified percentile range. The elasticity estimate is equal to 4.02 at the 33rd percentile (p33) and 5.39 at the 66th percentile (p66). All variables in the table are measured in logarithms except for **1**[·]. The table reports coefficient estimates from panel fixed-effect regressions. All specifications include both country and time fixed effects. Driscoll-Kraay standard errors are reported in parentheses. ***, **, and + indicate statistical significance at the 1%, 5%, and 15% levels, respectively. Observations are averages over (non-overlapping) five-year periods.

Table 5
Real exchange rate misalignments and manufactured exports.

Dependent variable	Manufactured exports		
Equilibrium *RER* calculation method	Balassa−Samuelson	H-P filter	Fundamentals
	(1)	(2)	(3)
RER overvaluation	-0.945***	-0.595**	-0.451**
	(0.093)	(0.181)	(0.157)
Observations	122	133	116
R^2 (within)	0.530	0.372	0.278
	(1)	(2)	(3)
RER undervaluation	-0.518	0.571*	0.159
	(0.400)	(0.253)	(0.344)
Observations	131	122	113
R^2 (within)	0.255	0.225	0.180

Notes: *RER* = real exchange rate. The table reports coefficient estimates from panel fixed-effect regressions. All specifications include both country and time fixed effects. Driscoll-Kraay standard errors are reported in parentheses. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. Observations are averages over (non-overlapping) five-year periods.

Table 6
The impact of commodity prices and different types of capital controls.

Estimation method	DOLS(1,1)					Panel FE				
Dependent variable	Real exchange rate					Manufactured exports				
Type of capital controls	FDI overall (1)	FDI inward (2)	Equity overall (3)	Equity inward (4)	GEQ inward (5)	FDI overall (6)	FDI inward (7)	Equity overall (8)	Equity inward (9)	GEQ inward (10)
RCP	0.685**	0.698**	0.155	0.112	0.697**	-2.009*	-2.024*	1.121	1.201	-2.017*
	(0.271)	(0.267)	(0.301)	(0.282)	(0.270)	(0.881)	(0.867)	(1.081)	(1.031)	(0.870)
RCP × *KC*	-0.010***	-0.010***	-0.000*	-0.000	-0.010***	0.008**	0.008**	0.000**	0.000*	0.008**
	(0.003)	(0.002)	(0.000)	(0.000)	(0.002)	(0.002)	(0.002)	(0.000)	(0.000)	(0.003)
KC	0.008**	0.008***	0.000	0.000	0.009***	-0.007**	-0.008**	-0.000**	-0.000*	-0.009**
	(0.003)	(0.003)	(0.000)	(0.000)	(0.003)	(0.003)	(0.002)	(0.000)	(0.000)	(0.003)
Government spending	0.222***	0.221***	0.263**	0.227*	0.222***					
	(0.061)	(0.060)	(0.115)	(0.133)	(0.060)					
Relative GDP per capita	2.101***	2.098***	0.256	0.383	2.118***					
	(0.374)	(0.378)	(0.411)	(0.386)	(0.384)					
Trade openness	-0.469***	-0.467***	-0.343***	-0.345***	-0.467***	1.127***	1.122***	0.833***	0.785***	1.118***
	(0.129)	(0.133)	(0.058)	(0.068)	(0.131)	(0.113)	(0.112)	(0.130)	(0.171)	(0.111)
Investment						0.354**	0.343**	0.278	0.324	0.336**
						(0.114)	(0.113)	(0.171)	(0.206)	(0.115)
Foreign income						0.051	0.049	0.073	0.054	0.046
						(0.073)	(0.072)	(0.196)	(0.170)	(0.072)
Observations	983	983	687	630	983	268	268	205	191	268
R^2 (within)	0.442	0.437	0.330	0.343	0.437	0.430	0.430	0.504	0.511	0.430
p-values for joint significance										
RCP and *RCP* × *KC*	< 0.01	< 0.01	< 0.01	0.445	< 0.01	0.012	0.011	0.026	0.075	0.014
KC and *RCP* × *KC*	< 0.01	< 0.01	< 0.01	0.053	< 0.01	0.044	0.026	0.019	0.134	0.047

Notes: *RCP* = real commodity price; *KC* = capital control. GEQ encompasses both FDI and portfolio equity. All variables in the table are measured in logarithms except for *KC*. DOLS(1,1) procedure includes contemporaneous, 1 lead, and 1 lag of changes of all continuous regressors, but they are suppressed to save space. All specifications include both country and time fixed effects. Driscoll-Kraay standard errors are reported in parentheses. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. The estimations are performed based on annual observations in columns (1)−(5) and non-overlapping five-year averages in columns (6)−(10).

Appendix A. List of sample countries and commodities

A.1. Sample countries

Our sample countries include: Argentina, Bangladesh, Bolivia, Brazil, Burundi, Central African Republic, Chile, Costa Rica, Cote d'Ivoire, Ghana, Guatemala, Honduras, India, Kenya, Madagascar, Malawi, Mali, Mauritania, Mauritius, Morocco, Mozambique, Niger, Pakistan, Papua New Guinea, Paraguay, Peru, Philippines, Senegal, South Africa, Sri Lanka, Sudan, Thailand, Togo, Turkey, Uganda, Uruguay, and Zambia.

A.2. Commodities used in the construction of commodity price indices

The list of 58 commodities includes: aluminum, bananas, barley, beef, coal, cocoa, coconut oil, coffee, copper, copra, cotton, crude oil (petroleum), fish, fishmeal, gold, groundnuts (peanuts), groundnut oil, hard logs, hardwood sawn, hides, iron ore, lamb, lead, maize, natural gas, nickel, olive oil, oranges, palm kernel oil, palm oil, phosphate rock, potash, poultry (chicken), rapeseed oil, rice, rubber, shrimp, silver, soft logs, softwood sawn, sorghum, soybean meal, soybean oil, soybeans, sugar, sunflower oil, swine, tea, tin, tobacco, TSP (triple superphosphate), uranium, urea, wheat, wood pulp, wool (coarse), wool (fine), and zinc.

Appendix B. Descriptive statistics

Table B.1
Summary statistics for variables used in the baseline regressions.

Variable	Frequency	Observations	Mean	Std. Dev.	Min	Max
RER	Annual	1,213	4.661	0.318	2.958	7.888
RER	5-year averages	269	4.674	0.294	3.919	7.146
RCP	Annual	1,406	1.025	0.390	0.094	2.340
RCP	5-year averages	296	1.026	0.390	0.099	2.308
KC	Annual	1,319	13.630	19.530	0.016	164.969
KC	5-year averages	294	12.214	18.002	0.017	123.045
Trade openness	Annual	1,378	-0.655	0.514	-2.761	0.341
Trade openness	5-year averages	291	-0.652	0.501	-2.124	0.293
Government spending	Annual	1,319	-2.079	0.374	-3.515	-0.297
Relative GDP per capita	Annual	1,384	0.779	0.123	0.476	1.174
Manufactured exports	5-year averages	276	-3.466	1.601	-12.951	-0.822
Investment	5-year averages	286	-1.632	0.350	-2.980	-0.680
Foreign income	5-year averages	293	9.479	0.810	6.613	11.771

Notes: *RER* = real exchange rate; *RCP* = real commodity price; *KC* = capital control.

Appendix C. Panel unit root and cointegration test results

Table C.1
Panel unit root tests.

	LLC		IPS		Fisher-ADF	
Variables	Level	1st diff.	Level	1st diff.	Level	1st diff.
RER	-0.25	-13.50***	2.30	-16.30***	2.82	-14.29***
RCP	-1.06	-7.78***	-0.38	-11.21***	-0.30	-10.18***
KC	-6.80***	-8.28***	-0.41	-15.14***	0.17	-12.19***
Trade openness	-1.63*	-18.26***	0.86	-18.32***	1.23	-15.56***
Government spending	-1.16	-17.71***	-1.01	-19.41***	-0.70	-16.70***
Relative GDP per capita	-5.20***	-4.86***	5.21	-11.44***	5.79	-7.94***

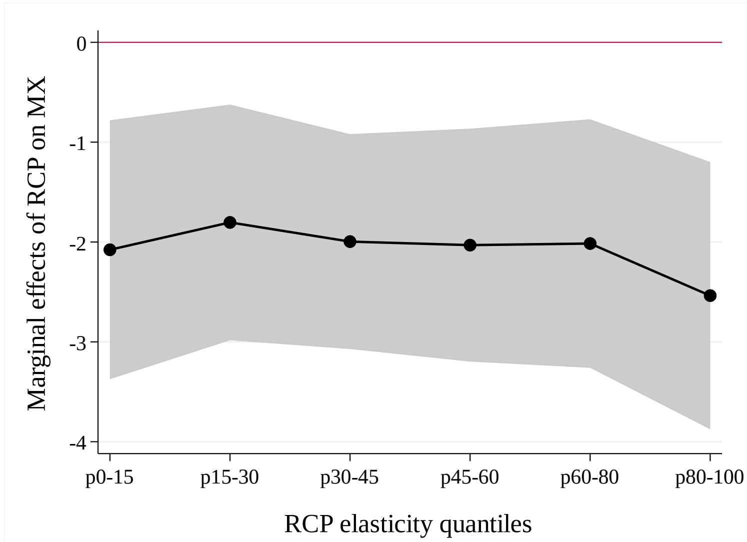
Notes: *RER* = real exchange rate; *RCP* = real commodity price; *KC* = capital control. The null hypothesis is that all panels contain a unit root. Reported are the t^*-statistics for the LLC test (Levin et al., 2002), W-statistics for the IPS test (Im et al., 2003), and Z-statistics for the Fisher-ADF test (Choi, 2001). For the series in levels, we include individual intercepts only in the LLC test, while both individual intercepts and trends are included in the IPS and Fisher-ADF tests. For the series in first differences, we include country-specific intercepts only for all tests. Lag lengths are automatically selected based on the modified Akaike information criterion. *** and * indicate the rejection of the null hypothesis at the 1% and 10% significance levels, respectively.

Table C.2
Panel cointegration tests.

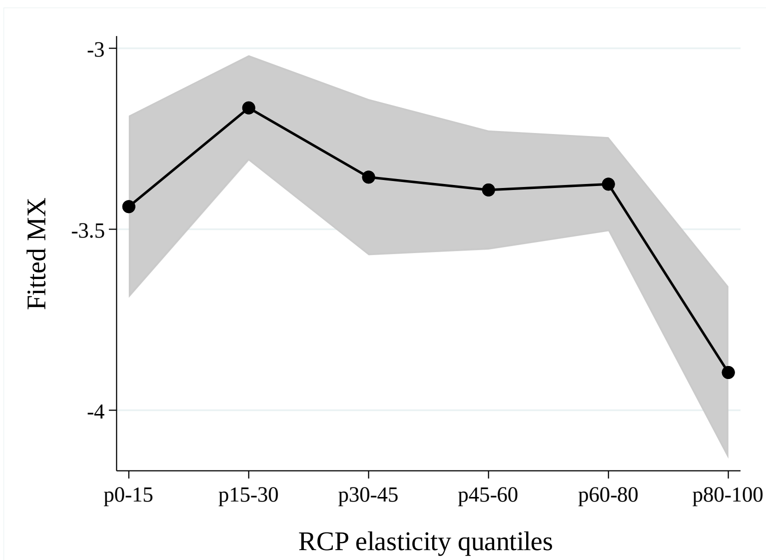
	Variables		
	RER, *RCP*	*RER*, *RCP*, *TO*, *G*, *YD*	*RER*, *RCP*, *TO*, *G*, *YD*, *KC*
Kao			
ADF t-statistic	-4.50***	-3.72***	-4.97***
Pedroni			
Within-dimension			
Panel v-statistic	-0.87	1.29*	-0.08
Panel ρ-statistic	-1.86**	-0.72	0.93
Panel PP-statistic	-4.04***	-8.76***	-7.92***
Panel ADF-statistic	-3.08***	-8.89***	-3.94***
Between-dimension			
Group ρ-statistic	2.09	2.71	5.27
Group PP-statistic	-0.39	-7.44***	-6.70***
Group ADF-statistic	-2.06**	-5.10***	-4.43***
Westerlund			
G_t	-2.30***		
G_a	-8.48*		
P_t	-17.06***		
P_a	-9.05***		

Notes: *RER* = real exchange rate; *RCP* = real commodity price; *TO* = trade openness; *G* = government spending; *YD* = relative GDP per capita; *KC* = capital control. The null hypothesis is that there is no cointegration. The Kao's (1999) cointegration test considers individual intercepts only, while the Pedroni's (2004) and Westerlund's (2007) tests allow for both individual intercepts and trends. Lag lengths are automatically selected based on the Akaike information criterion. ***, **, and * indicate the rejection of the null hypothesis at the 1%, 5%, and 10% significance levels, respectively. We were unable to perform the Westerlund's test for specifications in the last two columns of the table because of the insufficient number of continuous observations for some macroeconomic variables in low-income countries.

Appendix D. Supplementary results



(a) Marginal effects of commodity prices on manufactured exports



(b) Predicted manufactured exports

Fig. D.1. Marginal effects of commodity prices and predicted manufactured exports with other control variables.

Notes: *RCP* = (log) real commodity price; *MX* = (log) manufactured exports/GDP. In panel (a), we plot the marginal effects of commodity prices, $\partial MX / \partial RCP$, based on the model $MX_{it} = \sum_{k=1} \delta_k (RCP_{it} \times \mathbf{1}[\cdot]) + \boldsymbol{Y_{it}\gamma} + \phi_i + \phi_t + u_{it}$, where the indicator function $\mathbf{1}[\cdot]$ takes the value of one when the commodity price elasticity of real exchange rate for country *i* at time *t* falls within a specified percentile range; and $\boldsymbol{Y_{it}}$ is a vector of control variables, including *trade openness*, *investment*, and *foreign income*. Using the same model, panel (b) illustrates the predicted or fitted values of *MX* when all other variables including *RCP* take their sample average. The gray bands in both graphs represent 90% confidence intervals.

Table D.1
The impact of commodity prices and outward capital controls.

Estimation method	DOLS(1,1)			Panel FE		
Dependent variable	Real exchange rate			Manufactured exports		
Type of capital controls	FDI outward	Equity outward	GEQ outward	FDI outward	Equity outward	GEQ outward
	(1)	(2)	(3)	(4)	(5)	(6)
RCP	-0.614*	-0.314	-0.568*	-0.339	0.227	-0.666
	(0.302)	(0.415)	(0.312)	(1.079)	(1.242)	(1.289)
RCP × *KC*	-0.000	-0.000*	-0.000	-0.000*	-0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
KC	0.000	0.000+	0.000	0.000**	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Government spending	0.183	0.254*	0.217**			
	(0.109)	(0.128)	(0.104)			
Relative GDP per capita	1.963***	0.526	1.411***			
	(0.293)	(0.453)	(0.280)			
Trade openness	-0.330***	-0.414***	-0.333***	1.315***	0.791***	1.105***
	(0.060)	(0.058)	(0.052)	(0.125)	(0.149)	(0.157)
Investment				0.204*	0.280	0.255*
				(0.106)	(0.153)	(0.121)
Foreign income				0.065	-0.037	0.003
				(0.108)	(0.233)	(0.126)
Observations	770	537	823	228	170	239
R^2 (within)	0.383	0.402	0.351	0.327	0.520	0.327
p-values for joint significance						
RCP and *RCP* × *KC*	0.140	0.088	0.204	0.158	0.980	0.522
KC and *RCP* × *KC*	0.081	0.119	0.014	0.059	0.363	0.584

Notes: *RCP* = real commodity price; *KC* = capital control. GEQ encompasses both FDI and portfolio equity. All variables in the table are measured in logarithms except for *KC*. DOLS(1,1) procedure includes contemporaneous, 1 lead, and 1 lag of changes of all continuous regressors, but they are suppressed to save space. All specifications include both country and time fixed effects. Driscoll-Kraay standard errors are reported in parentheses. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. The estimations are performed based on annual observations in columns (1)−(3) and non-overlapping five-year averages in columns (4)−(6).

Appendix E. Model derivation for Section 2

E.1. Production

There are three sectors, *R*, *M*, and *N* with Cobb-Douglas technology:

$$Y_R = A_R L_R^{\alpha} K_R^{1-\alpha}, \tag{E.1}$$

$$Y_M = A_M L_M^{\beta} K_M^{1-\beta}, \tag{E.2}$$

$$Y_N = A_N L_N. \tag{E.3}$$

The zero profit conditions in each sector are given by

$$p_R A_R k_R^{1-\alpha} = r k_R + w, \tag{E.4}$$

$$p_M A_M k_M^{1-\beta} = r k_M + w, \tag{E.5}$$

$$p_N A_N = w, \tag{E.6}$$

where $k_i \equiv K_i / L_i$ is the capital-labor ratio in sector i. Log-differentiating the above conditions gives

$$\hat{p}_R = \mu_{L,R} \widehat{w} - \hat{A}_R, \tag{E.7}$$

$$\hat{p}_M = \mu_{L,M} \widehat{w} - \hat{A}_M, \tag{E.8}$$

$$\hat{p}_N = \widehat{w} - \hat{A}_N, \tag{E.9}$$

where $\mu_{L,i} \equiv wL_i/p_iY_i$ is the labor income share in sector i; and a hat above the variable denotes a logarithmic derivative, $\hat{x} = \mathrm{d}(\ln x)$. Assuming that $\hat{A}_R = \hat{A}_M$, we can combine Eqs. (E.7) and (E.8) to find the following wage rate response:

$$\hat{w} = \tau(\hat{p}_R - \hat{p}_M) \tag{E.10}$$

with $\tau = 1/\left(\mu_{L,R} - \mu_{L,M}\right)$.

Substituting Eq. (E.10) into (E.9), we can rewrite the price of nontraded goods as:

$$\hat{p}_N = \tau(\hat{p}_R - \hat{p}_M) - \hat{A}_N. \tag{E.11}$$

E.2. Real exchange rate

To find out the theoretical determinants of the real exchange rate in Eq. (E.12), we log-differentiate Eq. (10) in Section 2.3 and combine the result with Eq. (E.11) and the law of one price for the tradable goods:

$$\hat{Q} = \tau\theta\hat{p}_R^* - \theta^*\hat{p}_N^* + (\theta^* - (1+\tau)\theta)\hat{p}_M^* + \theta\hat{E} - \theta\hat{A}_N. \tag{E.12}$$

E.3. Capital controls

Deviating from the assumption of international capital mobility, we consider an extreme case of domestic capital market autarky. Now, the log-differentiated zero profit conditions in the two exportable sectors are given by:

$$\hat{p}_R = \mu_{L,R}\widehat{w} + \mu_{K,R}\hat{r} - \hat{A}_R, \tag{E.13}$$

$$\hat{p}_M = \mu_{L,M}\widehat{w} + \mu_{K,M}\hat{r} - \hat{A}_M, \tag{E.14}$$

where $\mu_{K,i} \equiv rK_i/p_iY_i$ is the capital income share in sector i. With a common rate of productivity shocks in the exportable sectors, combining Eqs. (E.13) and (E.14) generates the following capital return response:

$$\hat{r} = \frac{\hat{p}_R - \hat{p}_M - (\mu_{L,R} - \mu_{L,M})\widehat{w}}{\mu_{K,R} - \mu_{K,M}}. \tag{E.15}$$

Substituting Eq. (E.15) into (E.13) yields

$$\widehat{w} = \varphi\left(\hat{p}_R - \left(\frac{\mu_{K,R}}{\mu_{K,M}}\right)(\hat{p}_M + \hat{A}_R) + \hat{A}_R\right) \tag{E.16}$$

with $\varphi = 1/(\mu_{L,R} - (\mu_{K,R}/\mu_{K,M})\mu_{L,M})$.

Finally, substituting Eq. (E.16) into (E.9) gives

$$\hat{p}_N = \varphi\left(\hat{p}_R - \left(\frac{\mu_{K,R}}{\mu_{K,M}}\right)(\hat{p}_M + \hat{A}_R) + \hat{A}_R\right) - \hat{A}_N. \tag{E.17}$$