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Commodity currency reactions and the Dutch disease: the role of capital controls

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Abstract
Commodity booms generally induce real exchange rate appreciation in commodity-rich economies and make other tradable sectors less competitive. This "Dutch disease" phenomenon has been blamed for leading to structures of production with low diversification and undermining sustainable growth. In this paper, we explore whether capital controls can mitigate the transmission of commodity price changes to the real exchange rate and shield manufactured exports. Examining a panel of 37 commodity-abundant countries over the period 1980–2020, we find that a steeper commodity currency appreciation indeed has a more detrimental impact on manufactured exports. Restrictions on capital flows tend to reduce real appreciation pressures and the severity of the Dutch disease. Countercyclical capital controls seem to help foster economic diversification in commodity-dependent developing countries.

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

JEL Classification F14 · F3 · O13

1 Introduction

Commodity-rich economies often face large fluctuations in the value of their currencies due to the volatile global prices of their primary exports. These currency fluctuations can have detrimental impacts on the local economy. For example, persistent real appreciations could reduce competitiveness and investment in non-commodity export sectors. Conversely, sharp depreciations could increase the debt burden on domestic firms with large foreign liabilities. For these reasons, maintaining a competitive and

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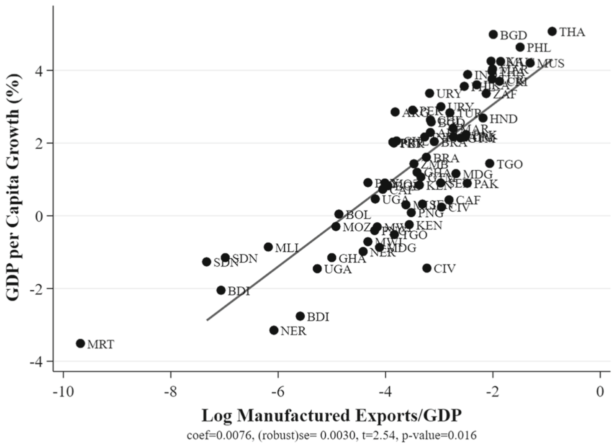

Fig. 1 Manufactured exports and GDP per capita growth, 1980 − 2020. *Note*: To obtain the fitted values, the growth rate of per capita GDP is regressed on manufactured exports, primary product (natural resource) exports, government spending, investment, trade openness, secondary schooling, population growth, and country and time fixed effects (all in logs except for the last three variables). Data source: World Bank's WDI

stable exchange rate may be of special interest to commodity-abundant developing countries.

In this paper, we focus on the role of countercyclical capital controls (i.e., tightening capital flow restrictions during booms and relaxing them during busts) in stabilizing the real exchange rate and preserving the competitiveness of manufactured exports in commodity-dependent developing economies.

To underscore the importance of manufactured exports to economic development, Fig. 1 displays relevant historical evidence in our sample of developing countries over the past four decades.[1] In the figure, each country has two observations for the log of manufactured exports as a ratio to GDP and the growth of GDP per capita, which are averages for each period, 1980 − 1999 and 2000 − 2020, so that we can trace their temporal changes within the economy. The illustration shows an apparent positive relationship between those two variables in our sample of commodity-exporting countries when controlling for other standard growth determinants. In line with this observation, Hausmann et al. (2007), Jones and Olken (2008), Berg et al. (2012), and Sheridan (2014) argue that growth accelerations are strongly associated with the development of the manufacturing export sector.[2]

The result in Fig. 1 suggests that commodity-rich developing countries may have an incentive to diversify their economies by expanding the manufacturing sector, which provides momentum for long-run economic growth. The primary purpose of this paper is to understand the challenges these countries face due to a significant reaction of the

[1] See Online Appendix A for a full list of sample countries.

[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).

real exchange rate to their commodity export prices and its effects on the long-run development of an industrial export sector.

We first study 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 nontradable 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. A stronger real exchange rate response is expected in countries more open to international capital transactions.

The theoretical framework generates two testable hypotheses: First, capital controls mitigate the transmission of commodity price changes to the real exchange rate. Second, capital controls reduce the propensity to crowd out manufactured exports resulting from a commodity price boom through muted exchange rate responses.

To explicitly test these hypotheses, we undertake a systematic panel data analysis based on a sample of 37 commodity-exporting developing countries over the 1980–2020 period. As a first step, we construct a country-specific real commodity export price index using the export volumes of 57 primary commodities and their global prices. 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 document statistically significant evidence that capital controls 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 a competitiveness-preserving role of targeted capital controls, we also report that the more excessive the commodity currency appreciation or real overvaluation, the worse the export performance of manufacturing. These results highlight the importance of countercyclical capital controls in countries whose currency values are strongly tied to commodity export prices to lower the intensity of the Dutch disease.[3]

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 a high cost of capital with limited borrowing sources. FDI inflows can bring 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.

However, our findings indicate that extreme caution needs to be used when opening up the capital market in countries specializing in commodity exports. Their real

[3] The Dutch disease refers to the coexistence of booming resource sector and lagging non-resource tradable sector due to increased input prices and currency appreciation. This is one of the classical theoretical justifications for the "resource curse," stagnant economic growth in resource-abundant developing countries despite their 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 theoretical developments, see Corden and Neary (1982), van Wijnbergen (1984), Krugman (1987), Matsuyama (1992), van der Ploeg and Venables (2013), and Alberola and Benigno (2017), among others. For empirical evidence related to this topic, see Ismail (2010), Bjørnland and Thorsrud (2016), Harding and Venables (2016), and Allcott and Keniston (2018).

exchange rates become more sensitive to commodity price cycles with a more open capital account. In these countries, countercyclical capital controls may implicitly subsidize economic diversification by stabilizing real exchange rate movements.[4,5]

In the next section, we review some related literature while emphasizing the paper's main contributions. Section 3 presents a simple small open economy model and derives two testable hypotheses. Section 4 describes the data and empirical model specifications. The baseline estimation results and robustness analyses are reported in Sect. 5, and finally, Sect. 6 concludes.

2 Related literature and contribution

This paper contributes to a vast literature on the Dutch disease and real exchange rate in three ways. First, we disentangle the mechanism of the Dutch disease into two key links—one from commodity prices to real exchange rates and the other from real exchange rates to manufactured exports—and jointly address them. These relationships have typically been studied separately in the prior literature. For example, the first link has been analyzed in the commodity currency literature by Chen and Rogoff (2003), Cashin et al. (2004), 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 harmful effects of real exchange rate uncertainty or misalignment. Unlike these studies, we examine the impacts of commodity price changes 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. Empirical evidence from Bodart et al. (2015) shows that an increase in commodity prices is related to stronger real appreciation when a country has a less open capital account. By contrast, Erten and Ocampo (2016) find that capital account regulations help decrease a real appreciation in emerging economies. Similarly, some studies report 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. Using such external shocks provides clear identification advantages in the empirical models of the Dutch disease. This argument can be justified by the

[4] 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.

[5] For capital controls and their role as a macroprudential policy, see the recent surveys provided in Engel (2016) and Erten et al. (2021).

notion that the world commodity price changes are driven mostly by global supply and demand conditions and can serve as an important source of an exogenous terms-of-trade shock to most 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., Lartey et al. 2012) or foreign aid flows and economic growth (e.g., Rajan and Subramanian 2008), it is less likely that our models suffer from a potential endogeneity bias.[6]

3 Theoretical framework and hypotheses

This section presents a minimal framework that stresses the transmission of commodity price changes to the real exchange rate and the resulting response in exports of manufactured goods. We build on the canonical small open economy model of Obstfeld and Rogoff (1996), with relevant implications taken from Bodart et al. (2015).

For our purpose, we assume that all of the commodity goods produced by the home country are exported abroad. The foreign country in the model is a home country's trading partner for manufactured goods but not commodity goods.[7] As is standard in the literature, we let global commodity prices be exogenously given to the domestic commodity sector.

3.1 Production and consumption

Consider that the domestic economy produces three types of goods: exportable commodities or resources (R), exportable manufactured goods (M), and labor-intensive 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, with α and β capturing the labor share.

In the benchmark case with the perfect international capital mobility, the domestic marginal product of capital is given by the world interest rate r^*, while perfect domestic

[6] 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).

[7] In other words, the rest of the world includes other countries trading the primary commodity products.

labor mobility ensures that the wage rate w is equalized across sectors. For simplicity, we assume a common rate of productivity shocks in the exportable sectors.

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

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

where $\tau = 1/\left(\mu_{L,R} - \mu_{L,M}\right)$; 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, $\hat{x} = \mathrm{d}(\ln x)$. As long as the commodity sector is more labor-intensive than manufacturing, we have $\tau > 0$.[8]

A domestic consumer's utility function takes the following form:

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

where C_N and C_M are the consumption of the two goods, θ is the share of nontraded goods consumption, and $\gamma = \theta^{-\theta}(1-\theta)^{-(1-\theta)}$. Similarly, the representative foreign household consumes the nontraded goods and imported manufactured goods produced by the home country.

3.2 Real exchange rate

Using the consumption-based aggregate price indices and the law of one price for tradable goods (i.e., $Ep_i = p_i^*$ for $i = M, R$), we can express the real exchange rate Q as follows:

$$Q = \frac{EP}{P^*} = \frac{Ep_N^{\theta} p_M^{1-\theta}}{\left(p_N^*\right)^{\theta^*}\left(p_M^*\right)^{1-\theta^*}}, \tag{6}$$

where E is the nominal exchange rate, defined as the price of domestic currency in terms of foreign currency; P is an aggregate price index; and a superscript asterisk on the variable denotes a foreign value. By construction, an increase in Q indicates a real appreciation of the home currency relative to the foreign currency.

3.3 Model implications and hypotheses

The parsimonious model structure enables us to derive three propositions, which form the basis of our main hypotheses:

Proposition 1 An increase in global commodity prices induces real appreciation in a commodity-exporting country: $\partial \hat{Q}/\partial \hat{p}_R^* > 0$.

[8] This assumption helps replicate the main logic of the resource movement effect, as in Corden and Neary (1982).

Proof By log-differentiating Eq. (6) 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 an increase in global commodity prices on the real exchange rate:

$$\frac{\partial \hat{Q}}{\partial \hat{p}_R^*} = \tau\theta > 0. \tag{7}$$

Given that labor is perfectly mobile between sectors and the price in the manufacturing industry 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.

Proposition 2 A commodity price boom crowds out manufactured exports through real appreciation: $\partial \widehat{C}_M^* / \partial \widehat{p}_R^* \leq 0$.

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

$$X_M = X_M\left(\frac{p_M}{P}\right), \tag{8}$$

$$C_M^* = C_M^*\left(\frac{p_M^*}{P^*}\right), \tag{9}$$

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 the tradable manufacturing sector and the definition of the real exchange rate, we find

$$\widehat{\left(\frac{p_M^*}{P^*}\right)} = \eta \hat{Q}, \tag{10}$$

where $\eta = \varepsilon^s / \left(\varepsilon^s - \varepsilon^d\right)$, $\varepsilon^s (\geq 0)$ 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. (10) 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. (9) with (10), we can derive Eq. (11), which demonstrates a decline in the home country's manufactured exports in response to rising global commodity prices, with the size of loss positively associated with the degree of real appreciation:

$$\frac{\partial \hat{C}_M^*}{\partial \hat{p}_R^*} = \varepsilon^d \eta \left(\frac{\partial \hat{Q}}{\partial \hat{p}_R^*}\right) \leq 0, \tag{11}$$

where $\partial\hat{Q}/\partial\hat{p}_R^* > 0$ by Eq. (7).

A surge in commodity prices tends to increase the domestic input costs (i.e., wage rates) of producing manufactured goods, appreciate the real exchange rate, and raise the relative price of manufactured goods for foreign consumers with an adverse effect on the foreign demand.

Proposition 3 Capital openness amplifies the magnitude of the exchange rate response to a commodity price change: $\left(\frac{\partial\hat{Q}}{\partial\hat{p}_R^*}\middle|\text{capital controls}\right) < \left(\frac{\partial\hat{Q}}{\partial\hat{p}_R^*}\middle|\text{open capital market}\right)$.

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 market. Resolving the model with only domestically mobile capital and labor, we find the following real exchange rate response to a commodity price change:

$$\frac{\partial\hat{Q}}{\partial\hat{p}_R^*} = \varphi\theta > 0 \tag{12}$$

where $\varphi = 1/\big(\mu_{L,R} - \big(\mu_{K,R}/\mu_{K,M}\big)\mu_{L,M}\big)$ with $\mu_{K,i}$ representing the capital income share $\big(0 < \mu_{K,i} < 1\big)$, defined as $\mu_{K,i} \equiv rK_i/p_iY_i$ in sector i. By comparing Eqs. (7) and (12), we observe that the real exchange rate reaction is smaller in the presence of capital controls because $\varphi < \tau$.

This proposition emerges 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 lower than would be the case with free international capital mobility. Consequently, the price of nontraded goods will increase less under capital controls, mitigating the appreciation pressures of the real exchange rate.

Combining propositions 1 and 3 gives the first testable hypothesis:

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

Moreover, combining propositions 2 and 3 generates the second testable hypothesis:

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

4 Data and empirical model specification

Our sample covers an unbalanced panel of 37 commodity-exporting countries from 1980 to 2020. 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.[9]

In the following, we explain the definition and source of the variables used in our empirical analysis and lay out the baseline regression models.

4.1 Key variables

4.1.1 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. It is measured such that a higher index corresponds to a real appreciation of the domestic currency. The data come from the Bruegel database released by Darvas (2012).

4.1.2 Real commodity price

Following Cashin et al. (2004), 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 exports. Specifically, we construct a country-specific real commodity price index using 57 commodities as follows[10]:

$$\mathrm{RCP_{it}} = \left[\sum_{j=1}^{J} w_{ij}\left(\ln p_{jt}\right)\right]/\mathrm{MUV}_t \tag{13}$$

where $w_{ij} = \left(1/T\sum_{t=1}^{T}\mathrm{ex}_{ij,t}\right)/\left(1/T\sum_{t=1}^{T}\mathrm{EX}_{it}\right)$; p_{jt} is the global price of commodity j at time t; MUV_t is the unit value index of manufactured exports for 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.

The global 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 commodity trade data from the UN COMTRADE database.

4.1.3 Capital controls

We construct a capital control variable based on an annual de facto international financial integration taken from the updated External Wealth of Nations Mark II database

[9] Large 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), complicating their economies' transmission mechanisms between resource export prices and real exchange rates. Moreover, almost all of them peg their currencies to the dollar and do not allow nominal exchange rate adjustments to an external shock.

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

(Lane and Milesi-Ferretti 2018). Among the integration indicators proposed by Lane and Milesi-Ferretti, we adopt the following measure of cross-border equity holdings:

$$\mathrm{GEQ}_{\mathrm{it}} = \left(\mathrm{EQ}_{\mathrm{it}}^{A} + \mathrm{FDI}_{\mathrm{it}}^{A} + \mathrm{EQ}_{\mathrm{it}}^{L} + \mathrm{FDI}_{\mathrm{it}}^{L}\right)/\mathrm{GDP}_{\mathrm{it}} \tag{14}$$

where $\mathrm{EQ}_{\mathrm{it}}$ and $\mathrm{FDI}_{\mathrm{it}}$ are, respectively, country i's stocks of portfolio equity and FDI at time t, with the superscript A indicating assets and the superscript L liabilities.

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. To create a capital control indicator, we take the inverse of GEQ so that a higher value of the indicator corresponds to stricter restrictions on capital flows.

4.1.4 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).

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

4.3 Baseline regression 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 Online Appendix Tables B.2 and B.3). 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.[11]

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

$$\begin{aligned}\mathrm{RER}_{\mathrm{it}} = \alpha_1 \mathrm{RCP}_{\mathrm{it}} + \alpha_2(\mathrm{RCP}_{\mathrm{it}} \times \mathrm{KC}_{\mathrm{it}}) + \alpha_3 \mathrm{KC}_{\mathrm{it}} + \boldsymbol{X}_{\mathrm{it}}\boldsymbol{\gamma} \\ + \sum_{j=-1}^{1} \Delta \boldsymbol{Z}_{\boldsymbol{i},\boldsymbol{t}+\boldsymbol{j}}\boldsymbol{\delta}_{\boldsymbol{j}} + \phi_i + \phi_t + \varepsilon_{\mathrm{it}}\end{aligned} \tag{15}$$

where $\mathrm{RER}_{\mathrm{it}}$ is the log of the real effective exchange rate; $\mathrm{RCP}_{\mathrm{it}}$ is the log of the real commodity price index; $\mathrm{KC}_{\mathrm{it}}$ is a capital control indicator; $\boldsymbol{X}_{\boldsymbol{it}}$ is a vector of additional fundamental determinants, including *government spending*, *relative GDP per capita*, and *trade openness*; $\boldsymbol{Z}_{\boldsymbol{it}}$ 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 a first difference operator. Controlling for the country and time fixed effects diminishes the problem of omitted variables bias or misspecification. We use Driscoll and Kraay's (1998) standard errors for statistical inferences to account for potential autocorrelation, cross-sectional correlation, and heteroscedasticity.

Our hypothesis 1 tests whether $\alpha_1 > 0$ and $\alpha_2 < 0$ in Eq. (15) so that the positive impact of *RCP* change 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; therefore, 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 (FE) estimator:

$$\mathrm{MX}_{\mathrm{it}} = \beta_1 \mathrm{RCP}_{\mathrm{it}} + \beta_2(\mathrm{RCP}_{\mathrm{it}} \times \mathrm{KC}_{\mathrm{it}}) + \beta_3 \mathrm{KC}_{\mathrm{it}} + \boldsymbol{Y}_{\boldsymbol{it}}\boldsymbol{\gamma} + \phi_i + \phi_t + e_{\mathrm{it}} \tag{16}$$

where MX_{it} is the log of the ratio of manufactured exports to GDP in country i at time t, and $\boldsymbol{Y}_{\boldsymbol{it}}$ is a vector of other potential determinants of country i's exports of manufacturing, including *trade openness*, *investment*, and *foreign income*. All other variables are as defined in Eq. (15).[12]

[11] 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." In fact, as reported in Table B.3, the Kao (1999), Pedroni (2004), and Westerlund (2005) test statistics mostly reject the null hypothesis of no cointegration, implying that the residuals are stationary and a set of variables are cointegrated. Note that the fully modified OLS (FMOLS), another widely used cointegration-estimation procedure, requires a balanced panel. When missing values in our sample are replaced using a linear extrapolation method, the FMOLS estimator produces similar results (available in Table B.4) to those generated by the DOLS.

[12] As displayed in Online Appendix Table B.5, the residual diagnostic test results indicate the presence of serial correlation and groupwise heteroscedasticity in the second baseline regression model in Eq. (16),

To focus on the long-run effects of commodity price movements on manufactured exports, we smooth out the business cycle effects by transforming our annual data into non-overlapping five-year averages, as is standard in the literature (e.g., Rodrik 2008).

Our hypothesis 2 tests whether $\beta_1 < 0$ and $\beta_2 > 0$ in Eq. (16) so that the negative impact of *RCP* change on *MX* may be moderated through restrictions on international capital movements. Regarding the other regressors, a greater value of the *investment* is likely to promote *MX* due to an increase in available physical capital, which may be required for manufacturing production. Higher *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.

5 Empirical results

5.1 Main results

The first three columns of Table 1 present the estimation results based on Eq. (15). The main parameters are the coefficients of the commodity price index *RCP* and its interaction with capital controls $RCP \times KC$.

Column (1) displays a significantly positive coefficient for *RCP*, demonstrating its long-run cointegrating relationship with *RER* in our sample countries. This result reinforces the previous empirical evidence for commodity currencies (Cashin et al. 2004).

Column (2) extends the specification with additional fundamental determinants of *RER*. We confirm a positive long-run relationship between *RCP* and *RER*, with expected signs for the other control variables. Indeed, including the other *RER* determinants strengthens the magnitude of *RCP* elasticity.

In column (3), significantly positive *RCP* and negative $RCP \times KC$ coefficient estimates indicate that while an increase in commodity prices induces real appreciation, a more stringent capital control appears to reduce the size of appreciation, supporting our hypothesis 1. In particular, a 1% rise in *RCP* would lead to a long-run real appreciation of 0.55% when *KC* is at its sample average and an appreciation of 0.43% when there is a one-standard-deviation increase in *KC* above its mean value.[13]

Turning to the *MX* regressions, we first show in column (4) a significantly negative response of *MX* to *RER* appreciation, consistent with the conventional theory. The

Footnote 12 continued
rationalizing our choice of the standard errors proposed by Driscoll and Kraay (1998). Further, according to the misspecification test results, there is little chance that the second baseline model is specified incorrectly and struggles with omitted variables.

[13] The net effects of a 1% increase in *RCP* are calculated by

$$\alpha_1 + (\alpha_2 \times \text{mean}_{KC})$$

and

$$\alpha_1 + (\alpha_2 \times (\text{mean}_{KC} + \sigma_{KC})),$$

respectively.

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	0.514*** (0.178)	0.685*** (0.245)	0.638*** (0.229)		− 2.068** (0.843)	− 1.748† (0.957)
RCP × *KC*			− 0.007** (0.003)			0.011** (0.004)
KC			0.006† (0.003)			− 0.010** (0.004)
Government spending		0.133† (0.079)	0.122† (0.079)			
Relative GDP per capita		1.544*** (0.359)	1.729*** (0.378)			
Trade openness		− 0.364*** (0.129)	− 0.363*** (0.115)		1.091*** (0.103)	1.084*** (0.107)
Investment					0.375** (0.125)	0.417** (0.129)
Foreign income					0.032 (0.068)	0.025 (0.065)
RER				− 0.640*** (0.167)		
Observations	1,265	1,156	1,147	255	268	268
R^2 (within)	0.160	0.306	0.356	0.266	0.402	0.408
p-values for joint significance						
RCP and *RCP* × *KC*			< 0.01			0.011
KC and *RCP* × *KC*			< 0.01			0.059

DOLS(1,1) procedure includes contemporaneous, one lead, and one 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)–(3) and non-overlapping five-year averages in columns (4)–(6)

negative coefficient estimate of *RER* indicates that a 1% increase in *RER* tends to lower *MX* by 0.64% in our sample countries.[14]

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. Other control variables such as *trade openness*, *investment*, and *foreign income* have the expected positive signs.

Finally, in column (6), we have a full specification, as in Eq. (16). A negative *RCP* coefficient and a positive coefficient for the interaction term lend support to our hypothesis 2. Specifically, a 1% rise in *RCP* would decrease *MX* by 1.61% when *KC* is at its sample average and by 1.41% 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 reveals that *KC* itself has a negative effect on *MX*. The reason may be that higher barriers to capital mobility can contract manufacturing production through a limited supply of inputs in the foreign capital-dependent assembly process or through foregone opportunities to benefit from positive spillovers generated by the commodity-sector FDIs. While the net effect of tighter *KC* on *MX* is positive in our sample, the negative standalone impact of *KC* suggests that a careful cost–benefit analysis across industries may precede the imposition of *KC* to exploit foreign capital more effectively.[15]

The last two rows of Table 1 report *p*-values for *F*-statistics to perform a joint significance test for *RCP* (or *KC*) and the interaction term. The consistently low *p*-values reported in columns (3) and (6) validate our baseline empirical specifications.

Since 18 countries in our sample specialize in agricultural food exports accounting for more than 50 percent of their total commodity exports on average, and eight countries in metal exports, we also explore the potential heterogeneity of the results across different types of commodities.[16] Interestingly, we find that the effects of *RCP* and its interaction with *KC* are less precisely estimated in the sample of agricultural food exporters and are not significantly related to *RER* and *MX*.

For the sample of metal exporters, we again find insignificant results in the *RER* regressions but statistically and economically significant evidence consistent with our baseline results in the *MX* regressions. Although we cannot draw a firm conclusion based on this small sample evidence, it is worth noting that metal exporters appear more vulnerable to the Dutch disease than agricultural food exporters. Capital controls

[14] 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.

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

[16] Costa Rica, Honduras, and Kenya are some examples of agricultural food exporters, and Chile, Mauritania, and Niger are some examples of metal exporters.

seem to be more capable of mitigating the extent of the Dutch disease in the metal-abundant economies as well. These results, available upon request, also suggest that it is important to consider a large panel of various commodity producers and their export price indices to identify commodity currency reactions and the role of capital controls on the severity of the Dutch disease.

5.2 Alternative capital control indicators

In this subsection, we test whether our main results are sensitive to alternative measures of capital controls. First, we use Chinn and Ito's (2006) index, 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 documented in the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions. As higher index values 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.[17]

Second, we employ the KOF de jure financial globalization index, available at the KOF Swiss Economic Institute (Gygli et al. 2019).[18] The de jure index is based on the capital account openness indicator developed by Chinn and Ito (2006) and the investment restrictions published in the World Economic Forum Global Competitiveness Report. It also reflects the number of international investment agreements to consider policies potentially encouraging capital flows. Since a higher index value represents that an economy is more financially globalized, we use the inverse of the KOF de jure 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 de jure index in columns (2) and (4). Indeed, the interaction effect between RCP and KC retains the expected signs in all cases, though it is not always strongly significant.

5.3 Controlling for exchange rate regimes and financial crises

To test the robustness of the main results, we introduce two more factors into the baseline regressions, which might affect the transmission of RCP changes to RER and MX.

First is a country's choice of exchange rate regime. We define a *flexible regime* dummy variable using the fine classification code of Ilzetzki et al. (2019). This dummy takes a value of 1 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 five-year average data, we first take the average of classification codes and then generate a binary regime variable following the same rule.

[17] We have also considered the 15th and 30th percentiles as alternative thresholds and found very similar results.

[18] The original KOF globalization index was introduced by Dreher (2006).

Table 2 Robustness check: using alternative capital controls

Estimation method	DOLS(1,1)		Panel FE	
Dependent variable	Real exchange rate		Manufactured exports	
Source of capital controls	Chinn–Ito	KOF de jure	Chinn–Ito	KOF de jure
	(1)	(2)	(3)	(4)
RCP	0.631** (0.265)	0.977*** (0.262)	− 1.997** (0.839)	− 2.763*** (0.640)
RCP × *KC*	− 0.136** (0.055)	− 0.123* (0.070)	0.301* (0.145)	0.342† (0.208)
KC	− 0.025 (0.042)	0.076 (0.059)	− 0.162 (0.106)	− 0.208 (0.188)
Government spending	0.077 (0.074)	0.143* (0.083)		
Relative GDP per capita	1.199*** (0.370)	1.234*** (0.416)		
Trade openness	− 0.354** (0.131)	− 0.412*** (0.131)	1.117*** (0.107)	1.225*** (0.130)
Investment			0.371** (0.128)	0.360** (0.130)
Foreign income			0.011 (0.070)	0.030 (0.066)
Observations	1,156	1,156	268	268
R^2 (within)	0.340	0.333	0.407	0.423
p-values for joint significance				
RCP and *RCP* × *KC*	< 0.01	< 0.01	< 0.01	< 0.01
KC and *RCP* × *KC*	< 0.01	< 0.01	0.143	< 0.01

DOLS(1,1) procedure includes contemporaneous, one lead, and one 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%, 10%, 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)

By construction, the reference category is a de facto peg or preannounced horizontal band with margins of no larger than $\pm$ 2%.[19]

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. 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 the *flexible regime* and its interaction with *RCP* are included as additional controls. First, we continue to see the expected signs on the parameters of our interest, with strong significance for the $RCP \times KC$ variable. However, the inclusion of multiple interaction variables that may be highly correlated lessens the statistical significance of the estimates for some of the regressors.

Moreover, in column (1), we find a significantly negative sign for *RCP*'s interaction with the *flexible regime*. This result reflects that an adjustable 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. Its coefficient estimate in column (3) has a negative sign, although statistically insignificant.

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 and 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.[20] In the case of the annual data, the *crisis* is a dummy variable to control for a crisis year. It intends to capture severe financial market instability that can cause significant changes for our dependent variables. The information for the banking crisis years comes from Laeven and Valencia (2018).

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.

5.4 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, therefore, investigate the cases for under- and overvaluations of *RER* relative to its equilibrium levels and their possibly different impacts on *MX*. Two versions of *RER* misalignments are considered here.

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

[20] Extending the *crisis* variable by covering the year 2020 as the COVID-19 recession does not alter our results at all.

Table 3 Robustness check: 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	0.736*** (0.240)	0.558** (0.244)	− 1.491 (1.023)	− 1.496 (0.951)
RCP × *KC*	− 0.009*** (0.003)	− 0.009*** (0.003)	0.011** (0.004)	0.011** (0.004)
RCP × *Flexible regime*	− 0.189*** (0.056)		− 0.319 (0.303)	
RCP × *Crisis*		0.171 (0.120)		− 0.614 (0.514)
KC	0.007** (0.003)	0.006* (0.003)	− 0.011** (0.004)	− 0.011** (0.004)
Flexible regime	0.025 (0.040)		0.299 (0.261)	
Crisis		- 0.095 (0.091)		0.209 (0.468)
Government spending	0.125† (0.078)	0.140* (0.081)		
Relative GDP per capita	1.729*** (0.396)	1.788*** (0.381)		
Trade openness	− 0.350*** (0.120)	− 0.422*** (0.112)	1.068*** (0.100)	1.074*** (0.137)
Investment			0.398** (0.119)	0.353** (0.137)
Foreign income			0.024 (0.064)	0.025 (0.063)
Observations	1,147	1,115	268	268
R^2 (within)	0.370	0.399	0.409	0.417
p-values for joint significance				
RCP and *RCP* × *KC*	< 0.01	< 0.01	0.013	< 0.01
KC and *RCP* × *KC*	< 0.01	< 0.01	0.070	0.056

DOLS(1,1) procedure includes contemporaneous, one lead, and one 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%, 10%, 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 4 Real exchange rate misalignments and manufactured exports

Dependent variable	Manufactured exports	
Equilibrium *RER* calculation method	Balassa–Samuelson (1)	Fundamentals (2)
RER overvaluation	− 0.043*** (0.003)	− 0.065*** (0.013)
Observations	121	109
R^2 (within)	0.520	0.472
RER undervaluation	− 0.009 (0.024)	0.026 (0.023)
Observations	132	146
R^2 (within)	0.212	0.141

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. *** indicates statistical significance at the 1% level. Observations are averages over (non-overlapping) five-year periods

The first approach follows Rodrik (2008) and defines 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. The second approach follows Goldfajn and Valdés (1999) and finds the predicted *RER* 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*.

Table 4 sets out the estimation results with *RER* overvaluation in the upper panel and undervaluation in the lower panel to see whether any asymmetry exists in *MX* responses. The upper panel of Table 4 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) and cointegrated fundamentals in column (2). These results are consistent with 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, meaning that *RER* undervaluation may not have a definite effect on manufactured exports. Overall, a central lesson is that excessive real appreciation is key to deterring export growth in the manufacturing sector.[21]

[21] We conduct an additional exercise by using the *RER* misalignment as a dependent variable in the estimation of our first baseline model in Eq. (15). We first test whether the *RER* misalignment series are nonstationary and cointegrated with a set of variables including *RCP*, and confirm the need for a cointegration estimator. As reported in Online Appendix Table B.6, the DOLS estimations show that our core results stay unchanged.

5.5 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 key driving forces behind the dampening role of capital controls, we disaggregate the *KC* variable into FDI versus portfolio equity and outward versus inward for each asset category.

We first generate the following financial integration indicators using the External Wealth of Nations dataset: FDI overall, FDI inward, FDI outward, (portfolio) equity overall, equity inward, equity outward, GEQ inward, and GEQ outward.[22] 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 measures. We then follow the procedure in Sect. 4.1.3 and create a proxy for capital controls by taking the inverse of the financial integration indicators for either direction for each category.

Table 5 summarizes the results when we redefine the *KC* variable at the disaggregated level with *RER* as the dependent variable in columns (1)−(5) and *MX* as the dependent variable in columns (6)−(10). We do not report the results with outward indicators, as the interaction variable of interest that involves them has almost no economic effects while statistically insignificant. This result aligns with our findings in Table 4 in that *RER* overvaluation is more of a concern than undervaluation, and overvaluation is more related to inward, rather than outward, capital movements.

Reviewing the results for *RER* regressions with FDI controls in 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 in columns (6) and (7).

When evaluating the net effects of an increase in *RCP* using the significant results in columns (4) and (9), we see that equity inflow regulations appear effective in stabilizing *RER* and insulating *MX* against rising commodity prices.[23] This is a notable finding because even if stock markets in our sample countries represent a relatively small fraction of their domestic economies, equity inflows in these countries seem to magnify *RER* appreciations and *MX* losses in the presence of a surge in *RCP*.

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

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

[23] Note that due to the missing observations for portfolio equity in some of our sample countries, regressions in columns (3) and (8) rely on 36 countries and those in (4) and (9) on 35 countries.

Table 5 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	FDI inward	Equity overall	Equity inward	GEQ inward	FDI overall	FDI inward	Equity overall	Equity inward	GEQ inward
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
RCP	0.627*** (0.229)	0.639*** (0.223)	0.172 (0.209)	0.458** (0.223)	0.637*** (0.227)	$-1.748^{\dagger}$ (0.954)	$-1.773^{\dagger}$ (0.937)	0.842 (0.733)	0.729 (0.772)	$-1.770^{\dagger}$ (0.941)
$RCP \times KC$	− 0.008** (0.003)	− 0.007*** (0.003)	− 0.00004 (0.00004)	− 0.0001*** (0.00003)	− 0.008*** (0.003)	0.011** (0.003)	0.010** (0.003)	0.0001 (0.0001)	0.0002*** (0.00004)	0.010** (0.003)
KC	0.006* (0.003)	0.005* (0.003)	0.00001 (0.00004)	0.00003 (0.00003)	0.006* (0.003)	− 0.009** (0.003)	− 0.010*** (0.003)	− 0.0001 (0.00005)	− 0.0001*** (0.00003)	− 0.011** (0.003)
Government spending	$0.122^{\dagger}$ (0.079)	$0.121^{\dagger}$ (0.079)	0.087 (0.124)	0.030 (0.111)	$0.121^{\dagger}$ (0.079)					
Relative GDP per capita	1.719*** (0.379)	1.719*** (0.376)	0.312 (0.410)	0.866* (0.448)	1.737*** (0.378)					

Table 5 (continued)

Estimation method	DOLS(1,1)					Panel FE				
Dependent variable	Real exchange rate					Manufactured exports				
Type of capital controls	FDI overall	FDI inward	Equity overall	Equity inward	GEQ inward	FDI overall	FDI inward	Equity overall	Equity inward	GEQ inward
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Trade openness	−0.367*** (0.117)	−0.367*** (0.119)	−0.247*** (0.055)	− 0.276*** (0.052)	− 0.366*** (0.119)	1.088*** (0.108)	1.080*** (0.106)	0.793*** (0.101)	0.817*** (0.139)	1.078*** (0.106)
Investment						0.423** (0.129)	0.409** (0.129)	0.378* (0.175)	0.391* (0.174)	0.404** (0.128)
Foreign income						0.027 (0.065)	0.024 (0.064)	0.060 (0.237)	0.166 (0.158)	0.022 (0.065)
Observations	1,147	1,147	842	776	1,147	268	268	209	195	268
R^2 (within)	0.359	0.356	0.245	0.296	0.354	0.408	0.407	0.505	0.531	0.407
p-values for joint significance										
RCP and *RCP* × *KC*	< 0.01	< 0.01	0.515	< 0.01	< 0.01	0.010	< 0.01	0.258	< 0.01	0.010
KC and *RCP* × *KC*	< 0.01	< 0.01	0.131	< 0.01	< 0.01	0.036	0.019	< 0.01	< 0.01	0.034

DOLS(1,1) procedure includes contemporaneous, one lead, and one 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%, 10%, and 15% levels, respectively. The estimations are performed based on annual observations in columns (1)−(5) and non-overlapping five-year averages in columns (6)−(10)

6 Conclusion

Slow economic growth in developing countries that rely heavily on raw commodity products has been a long-standing topic in economics. Indeed, the resource curse literature extensively documents that while commodity windfall gains have positive short-run impacts on economic growth, their long-term effects tend to be negative (Collier and Goderis 2012). Accordingly, 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 (Matsuyama 1992; Gylfason et al. 1999).

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

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 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 fostering export diversification in resource-rich developing countries.

Policies facilitating investments in infrastructure, education, and R&D can also encourage the production of the manufacturing sectors and complement targeted capital controls to further enhance the growth potential of the countries.

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Declarations

Conflict of interest The authors declare that they have no conflict of interest.

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