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Pricing-to-market of Brazilian exports: a case of vehicle currency invoicing

Sérgio Kannebley Júnior
Department of Economics, University of São Paulo, Ribeirão Preto, Brazil
Diogo de Prince
Department of Economics, Federal University of São Paulo, Osasco, Brazil, and
Daniel Quinaud Pedron da Silva
Department of Economics, University of São Paulo, Ribeirão Preto, Brazil

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Abstract

Purpose – Brazil uses the dollar as a vehicle currency to invoice its exports. This fact produces a tendency toward equalizing the prices of products in dollars in the international market and reducing the ability of firms to practice pricing-to-market (PTM). This study aims to evaluate the hypothesis by estimating error correction models in panel data, obtaining estimates of PTM for 25 manufacturing products exported by Brazil between 2010 and 2020.

Design/methodology/approach – This study uses the correlated common effect estimator proposed by Pesaran (2006) and Chudik and Pesaran (2015b) to estimate the PTM coefficients.

Findings – Results of this study indicate that exporters practice local-currency pricing stability for dollar prices. This study obtains that Brazilian exporters tend to stabilize their dollar price for exports, reducing heterogeneity between destination markets. The results are in agreement with the hypothesis of the prevalence of the coalescing effect of Goldberg and Tille (2008) and lower sensitivity of the markup adjustment to the specific market, as pointed out by Corsetti *et al.* (2018). The pricing of Brazilian exports in dollars reflects a profit maximization strategy that considers an international price system based on global demand for products.

Originality/value – In addition to analyzing the dollar role in the pricing of Brazilian exports through the triangular decomposition, this study also shows the importance of examining the cross-section dependence of errors, considering the heterogeneous cointegration in export pricing models and producing PTM estimates for short-term and long-term.

Keywords Heterogeneous cointegration, International trade, Pricing-to-market, Error correction models, Cross-section dependence

Paper type Research paper

1. Introduction

The United States dollar (USD) is used in international trade for invoicing transactions and has prominence in financial transactions around the world, serving as currency vehicle pricing (Goldberg, 2010; Goldberg & Tille, 2016) [1]. Goldberg (2010) lists reasons for invoicing exports in dollars even if the exporter or importer is not the USA, including the incentive for exporters to limit price movements relative competitors by choosing the invoicing currency used by the majority of producers in the industry, denominated as the “coalescing” effect [2].

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Strategic complementarities can be a reason for using vehicle currency for export pricing. Firms that face strategic complementarities keep stable prices compared to competitors seeking to maximize profits, according to Goldberg and Tille (2008). Brazil is a case of using the dollar as a vehicle currency in its foreign trade transactions. According to data in studies by Ilzetzki, Reinhart, and Rogoff (2019) and Boz *et al.* (2022), 94% of exports and 84% of Brazilian imports are invoiced in dollars between 2001 and 2017, although Brazil is not an exception in this regard [3]. The high macroeconomic volatility experienced by the country since the mid-twentieth century justified this practice, which led to the adoption of exchange rate arrangements pegged to the dollar. This case of the prevalence of the dollar as a vehicle currency makes it interesting to analyze the pricing-to-market (PTM) hypothesis for Brazil. PTM refers to a company selling the same good to multiple destination markets with divergent prices between them, given that exchange rate variations lead to different markup responses across locations (Burstein & Gopinath, 2014). This article seeks to test PTM in the presence of vehicle currency based on data from 25 products exported by Brazil to 18 destination markets between 2010 and 2020.

Firm decisions to set export prices and choose the currency for invoicing are interrelated in an environment with price rigidity. Correa, Petrassi, and Santos (2018) report that Brazilian firms show some rigidity degree in setting prices and that the exchange rate is the most important driver of price changes for exporting firms. According to these authors, around 58% of surveyed companies set prices in Brazilian real and then convert them into the currency of the foreign country. However, the authors are not confident in this answer by mentioning that this would suggest *"that most of Brazilian exporting firms either have some market power in international markets or answer without reporting actual practice"* (Correa *et al.*, 2018, p. 300) [4].

On the other hand, international evidence indicates the currency in which exports and imports are invoiced is a good proxy for the currency in which firms set prices (Corsetti, Crowley, and Han, 2018). Considering that most Brazilian manufacturing companies do not have a relevant share of the international market and use the dollar to invoice their exports, prices are more likely to be fixed in dollars. In this case, we should expect relative stability of Brazilian export prices in dollars compared to their international competitors. For this, markup variations in domestic currency should be negatively associated with the change in the exchange rate between the Brazilian currency and the dollar, with a reduced difference between the destination markets. In a strict sense, although this would be a case of variable markup, it would limit the practice of PTM as the markup responses were more homogeneous between the destination markets. There is evidence that PTM is usually related to the pricing of exports based on the currency used at the destination market (Gil-Pareja, 2003; Fitzgerald & Haller, 2014; Corsetti *et al.*, 2018). However, evidence for the practice of PTM related to the use of vehicle currency or producer currency is scarce (Corsetti *et al.*, 2018). We aim to test the PTM hypothesis for Brazilian exports by evaluating the hypothesis that prices are fixed in dollars.

Boz, Gopinath, and Plagborg-Møller (2019) obtain that variation across country pairs in exchange rate pass-through and trade elasticity is meaningfully explained by the dollar's dominance as invoicing currency, corroborating predictions of the "dominant currency paradigm" (DCP). Under the DCP, Gopinath *et al.* (2020) derive predictions for asymmetry from the effects that shocks to the dollar have on global trade, consumption and production. If there are monetary policy shocks in response to global or local shocks on the United States economy, then there would be an asymmetric transmission mechanism of the effects to national monetary policies. This should lead to a correlation between countries' exchange rates against the dollar. Still, this should imply a cross-section dependence (CD) between exchange rates due to their relationship with the US currency. It produces an endogenous relationship between prices in international trade and exchange rates due to this common factor arising from the dominance of the US dollar in international trade.

We evaluate the PTM hypothesis considering that the hypotheses of exogeneity of the exchange rate and independence of errors between destination markets in the traditional estimation for PTM based on Knetter (1989) may no longer hold. That is, trade flows, asset prices and inflation in countries other than the USA would be directly influenced by international movements in the value of the dollar considering the assumption that the US dollar is a dominant currency in the financial, goods and services markets (Goldberg, 2010).

We use unit value as a proxy for export price, where unit value is the ratio between export value and export quantity. We are aware that unit value data as a proxy for price data is not ideal [5]. Thus, we cannot associate the product information to the firm and, therefore, control for the heterogeneities of exporting firms. The usual solution is to assume the homogeneity of the exporters' marginal costs – or their evolution – between destination markets and represent them from a common homogeneous factor (Knetter, 1989).

In response to these two problems, we adopt a representation with unobserved common factors that help to mitigate these identification problems. This representation tends to correct problems arising from the presence of measurement errors by allowing heterogeneity in the response to common factors that are supposed to originate from cost variations and correct the CD of the errors. We use the common correlated effect pooled (CCEP) and common correlated effect mean group (CCEMG) estimators from Pesaran (2006) and Chudik and Pesaran (2015b). Also, we combine an error correction model (ECM) proposed by Gagnon and Knetter (1995) with Eberhardt and Presbitero (2015) modeling strategy for panel data. CCEMG is an econometric technique that controls for the common correlated effect (CCE), addresses the potential nonstationarity of the data with ECM specification and produces heterogeneous estimates of the parameters. Using panel data estimators, we test the PTM hypothesis for industrial products exported from Brazil and we produce estimates for the degree of PTM in the short and long run based on the Gagnon and Knetter (1995) model.

Considering the hypothesis that exporters price their products in dollars, we estimate two specifications that the only change between them is the exchange rate variable used as an explanatory variable. The first follows the usual form proposed by Knetter (1989) and Gagnon and Knetter (1995), which uses the bilateral exchange rate of the currency of the destination market in terms of the domestic currency. In the same spirit as Boz *et al.* (2022), which explores the role of vehicle currency invoicing for exchange rate pass-through, our second specification considers the triangular decomposition of the exchange rate in terms of the dollar [6]. Thus, we can compare the results from the different specifications, test the hypothesis that prices are effectively established in dollars and measure the markup adjustment against the different sources of exchange rate variation. Our contribution is twofold in that sense. In addition to analyzing the dollar role in the pricing of Brazilian exports through the triangular decomposition, we also show the importance of examining the CD of errors, considering the heterogeneous cointegration in export pricing models and producing PTM estimates for short-term and long-term.

Our results indicate that the CCEMG estimator reduces the CD of residuals with the error correction models, producing a statistically more adequate specification. Our results support the evidence that export prices are sensitive to exchange rate variation in two ways: (i) the destination market currency in relation to the domestic currency and (ii) the dollar in relation to the domestic currency. We obtain a heterogeneous long-term relation across countries in both specifications. A 10% of depreciation of the exporter's currency relative to a particular destination market leads to a 9.1% destination-specific reduction in the markup of export price over marginal cost.

However, in the second specification that uses the triangular decomposition of the exchange rate around the dollar, the estimated coefficients indicate a complete adjustment of prices in domestic currency to stabilize prices in dollars in the foreign market. We find that a 10% of depreciation of the exporter's currency relative to a particular destination market leads to a 10.5% reduction in the markup of this destination market. Furthermore, the variation in the exchange rate of the currency of the destination market in terms of dollars is not relevant to the pricing of Brazilian exports considered in the sample. The heterogeneity in the markup adjustment response across destination markets is greatly reduced when we estimate with this alternative specification. Our results are in agreement with the hypothesis of the prevalence of the coalescing effect of Goldberg and Tille (2008) and of lower sensitivity of the markup adjustment to the specific market as pointed out by Corsetti *et al.* (2018). Even though our sample is composed of differentiated goods according to the Rauch (1999) classification, which are the goods that have the lowest probability of predominance of the coalescing effect for exports according to Goldberg and Tille (2008), we obtain results indicating coalescing effect.

PTM estimates indicate that Brazilian exporters set their products' prices in USD on the international market for 24 of 25 products among the analyzed products. Under the assumption of an international price system in place, this would discourage discrimination between destination markets. The predominance of export prices set in US dollars makes exporters' profitability extremely dependent on variations in the Brazilian currency exchange rate in relation to the dollar. Thus, instead of domestic currency depreciation leading to a perception of cheaper Brazilian industrial exports, what we actually have is price stability in dollars and an increase in profit margins. However, our evidence suggests that there is no specific markup adjustment to the destination market. This would contradict the PTM hypothesis for Brazilian manufacturing exports from our analysis.

We divide the rest of this article into seven sections, in addition to this introduction (Section 1). Section 2 presents a review of the literature on PTM and currency invoice. Section 3 describes the empirical model to be estimated, highlighting the need for a representation with unobserved common factors. Section 4 presents the econometric methodology used in the article, and we report the descriptive statistics from the data in Section 5. In Section 6, we discuss the results of the dynamic models and we analyze the sources of heterogeneity of the coefficients according to the different specifications used. We seek to address the economic context of the results and the implications for policymakers in Section 7, along with the discussion of the literature. Lastly, we present the final comments in Section 8.

2. Pricing-to-market and currency choice

PTM consists of making markup adjustments for the same good in different destination markets because of exchange rate variation, producing a divergence in prices between different markets. More specifically, Burstein and Gopinath (2014) show that PTM requires not just variable markups but the response of markups should vary across locations.

Knetter (1995) provides the static representation, considering the problem of maximizing the profit of an exporter of a good to N different segmented markets with flexible prices [7]. The profit function is given by:

$$\Pi(p_1, p_2, \ldots, p_N) = \sum_{i=1}^{N} p_i q_i(E_i p_i, \nu_i) - C\left(\sum_{i=1}^{N} q_i(E_i p_i, \nu_i), \omega\right)\delta_t \tag{1}$$

where p is the price in the exporter's currency, q is the quantity demanded, which is a function of the price in the currency of the destination market i, $E_i p_i$, ν_i is a demand shock specific to the market, E_i is the nominal exchange rate (units of the currency of the destination market per unit of the exporter's currency), $C(q, \omega)$ is the cost function, where ω represents the input price and δ_t is a random shock that may shift the cost function. Profit maximization produces the following first-order condition:

$$p_i = MC_t\left(\frac{\eta_i}{\eta_i - 1}\right) \tag{2}$$

where $i = 1, \ldots, N, MC_t = \left(\frac{\partial C}{\partial q}\right)\delta_t$ is the marginal cost and η_i is the price elasticity of demand in market i. Thus, the export price to each destination market is the common marginal cost times the destination-specific markup. A change in the exchange rate vis-à-vis the currency of country i can affect the price charged in market i in two ways: by affecting marginal cost (through changes in quantity or input prices) and the price elasticity of demand. The former effect is common for all other destination markets, while the latter is destination-specific. Both effects determine exchange rate pass-through, while PTM refers only to the second effect. In this section, we focus on the sensitivity of the price elasticity of demand and the markup to exchange rate variations.

Krugman (1986) notes that the monopolist price discrimination model can only explain PTM if demand curves have the right shape. The general rule is the following. If demand, as perceived by the firm, becomes more (less) elastic as local currency prices rise, then the optimal markup charged by the exporter will fall (rise) as the buyer's currency depreciates. No matter what model of market structure we assume, the price adjustment in response to exchange rate changes depends on the firms' perceptions of how demand elasticities change with respect to the local currency price. Knetter (1995) demonstrates this argument by totally differentiating Equation (2), yielding the following expression:

$$\frac{dp_i}{p_i} = (1+\beta_i)\frac{dC_q}{C_q} + \beta_i\frac{dE_i}{E_i} \tag{3}$$

where $\beta_i = \frac{\frac{\delta \ln \eta_i}{\delta \ln p_i^*}}{\left((1-\eta_i) - \frac{\delta \ln \eta_i}{\delta \ln p_i^*}\right)}$, $p_i^* = E_i p_i$ is the price in the buyer's currency, and $\frac{dC_q}{C_q}$ equals to the total differential of the logarithm of marginal cost. A demand curve that is less convex than the constant elasticity of substitution (CES) should present $\frac{\delta \ln \eta_i}{\delta \ln p_i^*} > 0$, making the denominator of that fraction negative and, therefore, $\beta_i < 0$. This means that as the importer's currency depreciates, there is a price increase (in local currency) that will be partially offset by a reduction in the markup as the demand curve becomes more elastic. This PTM behavior is known as local currency price stability (LCPS). In turn, a more convex demand curve than the CES will lead to the opposite result, i.e. $\beta_i > 0$, with increasing markups as the exchange rate devaluates.

According to Knetter (1993), the discussion about the sign of β depends on hypotheses of market segmentation and the imposition of an additional structure on the models to ensure that the markup responds to the prices. Burstein and Gopinath (2014) discuss models with non-CES demand specifications based on Kimball (1995) – in which the markup elasticity is higher for low-relative-price firms – with strategic complementarities based on Atkeson and Burstein (2008) – where the markup elasticity is higher for higher-market-share firms – and with distribution costs based on Corsetti and Dedola (2005) – where the markup elasticity is higher for firms with higher distribution share. With strategic complementarities in pricing, exporters firms would want to keep their prices stable in relation to the prices of their competitors, adjusting their markups to avoid losing market share.

Gopinath (2015) concludes that sticky price concerns in international trade are well-founded based on evidence of price rigidity produced by microdata price studies. According to Burstein and Gopinath (2014), the firm can choose its currency to keep its preset price closer to the desired price when prices are sticky. Thus, in the presence of rigidity in prices, the currency in which the product is invoiced becomes relevant to measure the exchange rate pass-through and specifically the adjustment of the markup to the exchange rate variation.

The terminology is Producer Currency Pricing (PCP) when prices are rigid in the producer's currency. In this case, firms set export prices in domestic currency, and the foreign currency prices of its products vary with the exchange rate. Following Corsetti, Dedola, and Leduc (2010), the alternative view is that firms preset prices in domestic currency for the domestic market and in foreign currency for the destination market. This hypothesis is Local Currency Pricing (LCP). However, the growing evidence indicates that most global trade transactions are invoiced in just a few currencies regardless of the countries involved in the transaction. This strengthens a third possibility, that firms set a price on a vehicle currency (VCP).

The choice of currency to set prices and the adjustment of prices by exchange rate shocks depend on the expected profit maximization strategy that the firm intends to follow, given the frequency of price adjustment and the magnitude of the exchange rate pass-through that it intends to carry out while prices are not readjusted. According to Burstein and Gopinath (2014), the choice of invoicing currency depends on the unconditional desired pass-through, that is, the value of pass-through if the firm could change the price flexibly. In this sense, the discussion on the adjustment of the markup to exchange rate variations returns to the arguments that determine how the price elasticity of demand varies in response to exchange rate shocks in a flexible price environment.

If the firm wants a lower exchange rate pass-through, then it is better for the firm to set prices in consumer currency (LCP), ensuring a zero exchange rate pass-through in the importer's currency and consequently an adjustment in the exporter's markup. On the other hand, if the firm wishes to carry out a high exchange rate pass-through in the importer's currency in the short term, then it must choose to fix prices in the producer's currency, keeping the exporter's markup unchanged against an exchange rate variation. In the presence of strategic complementarities, the choice of currency to price exports may depend on the option of other competitors in the international market. In this sense, if foreign competitors price their products in vehicle currency, then the exporter is likely to price their products in the same currency (VCP). Corsetti *et al.* (2018) state that pricing in vehicle currency overcomes market segmentation and translates into a "reference price system," by which firms do not exploit market-specific demand elasticities, but price in relation to global demand. According to Goldberg and Tille (2008), the coalescing effect that is the use of vehicle currency occurs mainly in industries whose goods have a lower degree of heterogeneity and whose markets do not present a lower degree of concentration. Their evidence indicates that factors such as the size of the exporting country and the destination market are also positively associated with the use of vehicle currency.

Gil-Pareja (2003) argues that the combination of price rigidity and invoicing in the importer's currency can lead to spurious findings of PTM behavior. Gil-Pareja (2003) uses price data in the European car market between 1993 and 1998 to test whether the type of invoicing has a significant impact on PTM patterns. His findings indicate that LCPS is a strong and pervasive phenomenon across products independent of the invoicing currency. Fitzgerald and Haller (2014) analyze data from Irish exporters selling to the UK using micro data on Irish producer prices. They obtain that PTM is a practice by exporters. Thus, when prices are fixed in the local currency – conditional on price fluctuation –, producers choose markups such that the ratio between the markups in the foreign and domestic markets increases (declines) equiproportionally with the depreciation (appreciation) of the domestic currency. Corsetti *et al.* (2018) study if the invoicing choice is related to firms' strategic markup adjustments, i.e. pricing to market and their findings. They use micro evidence from UK customs transactions from 2010 to 2016 that includes information on the invoicing currency of a transaction. They find a substantial destination-specific markup adjustment only for export shipments that are invoiced in the destination market's currency, consistent with the view that firms set prices in domestic currency for the domestic market as the LCP hypothesis. Conversely, they do not obtain destination-specific markup adjustments by firms that invoice a shipment in either their own currency or a vehicle currency, consistent with a firm setting one price either in their own or in a vehicle currency. Later in the article, we will address how (i) exporting firms set prices in a third currency such as the dollar (vehicle currency) and (ii) the possibility of a coalescing effect with the sample of products considered can affect the results and the specification used according to the discussed literature.

3. Empirical model for pricing-to-market

From an empirical point of view, Goldberg and Knetter (1997) argue that estimating the degree of PTM involves problems of measurement error and endogeneity because firms' marginal costs are unobservable and by the effect of exchange rate fluctuations on production and the costs of imported inputs, which impact the marginal cost. Knetter (1989) develops an identification strategy, seeking to separate the effects of exchange rate fluctuations on cost and markup changes by estimating the following two-way fixed-effect model:

$$\tilde{p}_{it} = \psi_i + \lambda_t + \beta_i \tilde{E}_{it} + u_{it} \tag{4}$$

where $\tilde{p}_{it}$ corresponds to the logarithm of export price in units of the exporter's currency, ψ_i is a set of destination country effects, λ_t is a set of time effects, $\tilde{E}_{it}$ is the logarithm of the nominal exchange rate in the exporting country relative to the destination market, deflated by the consumer price index in the destination market, u_{it} is a random error, i refers to the destination country and t is the time period [8]. $\sim$ indicates the logarithm of the variable.

In Equation (4), the markup component is represented by the additive combination of a market-specific term, ψ_i and the term $\beta_i \tilde{E}_{it}$. According to Equation (3), in which β_i is the response of the price elasticity of demand to the variation of the price in foreign currency (of the destination market), this coefficient is only statistically significant under the hypothesis that the price elasticity of demand is sensitive to the exchange rate variation. Considering that the same plant exports to several locations, changes in marginal cost affect export prices to all destination markets equally. In this sense, we assume that homogeneous time effects are able to capture common cost changes across destinations following Knetter (1989).

The weak correlation between the general equilibrium effects on the demand side and the exchange rate justifies the exogeneity assumption of the exchange rate variable. This implies an absence of correlation between exchange rates and omitted variables that influence the elasticities in the different markets. However, this exogeneity assumption is no longer true when considering the possibility of global supply or demand shocks also affecting exchange rates.

Dollar is the dominant anchor currency, i.e. central banks aim to stabilize their own currency against dollar (Ilzetzki *et al.*, 2019). This role of the dollar as the dominant anchor currency and a share of trade being invoiced in dollar propagate US monetary policy impulses to other countries and provide a common component to the global monetary environment according to Gourinchas (2021). About this point, import price inflation in many countries should depend on movements in their exchange rate against the dollar regardless of the share of their trade with the USA (Gopinath & Itskhoki, 2022). Therefore, we can no longer consider that omitted factors are not correlated with the exchange rate. In this case, we should assume the presence of a common global factor present in the dynamics of exchange rates and in the inflationary process of economies.

In econometric terms, this translates to the presence of unobserved common factors (which would not homogeneously impact each observation unit) and a relationship between those factors and the model's explanatory variable(s). Given those potential problems for panel data estimators, we modify the specification of the empirical model in Knetter (1989) to capture this unobserved heterogeneity. We have the following general structure:

$$\tilde{p}_{it} = \beta_i \tilde{E}_{it} + u_{it} \tag{5}$$

$$u_{it} = \alpha_i + \lambda_i' f_t + \varepsilon_{it} \tag{6}$$

$$\tilde{E}_{it} = \pi_{i,E} + \gamma_i' f_{E,t} + v_{it} \tag{7}$$

$$f_t = \delta' f_{t-1} + \zeta_{it} \tag{8}$$

In Equation (6), α_i corresponds to country-specific effects, while f_t is a set of common factors with country-specific factor loadings λ_i. In Equation (7), $\tilde{E}_{it}$ is also affected by unobservable common effect, f_t, as $f_{E,t} \subset f_t$ has a set of common factors with country-specific factor loading γ_i. We assume that the error terms ε_{it}, v_{it} and ζ_{it} are random and not correlated. This setup leads to the endogeneity of $\tilde{E}_{it}$ in Equation (5), where the regressor $\tilde{E}_{it}$ is correlated with the unobservable u_{it} through the common factors.

Additionally, we do not assume stationarity of the variables $\tilde{p}_{it}$, $\tilde{E}_{it}$ or of the vector of unobservable common factors, f_t, allowing the presence of heterogeneous cointegration. Similar to what Knetter (1989) did, the fixed effects α_i should capture the different levels of markup specific to each destination market. However, unlike Knetter (1989), we do not impose the restriction of a representative homogeneous common factor for the evolution of the marginal cost and we do not restrict the existence of a single unobservable common factor as time dummy variables.

The presence of exchange rate fluctuations that are unanticipated – or expected to reverse – combined with supply or demand adjustment costs can generate differences in the degree of PTM in the short and long term (Krugman, 1986; Kasa, 1992; Gagnon & Knetter, 1995). Considering a dynamic process of adjustment to long-run equilibrium, we include a lagged dependent variable and/or lagged regressors in the econometric model. With nonstationary variables and the existence of a cointegrating relation between the variables, we use an error correction model with panel data. Thus, we distinguish the behavior of PTM between the short and long run, and we write ECM similarly to Gagnon and Knetter (1995) as:

$$\Delta\tilde{p}_{it} = \alpha_i + \rho_i \left(\tilde{p}_{it-1} - \beta_i \tilde{E}_{it-1} - \lambda_i f_{t-1} \right) + \gamma_{E,i} \Delta\tilde{E}_{it} + \gamma_{f,i} \Delta f_t + \varepsilon_{it} \tag{9}$$

where β_i and λ_i coefficients represent the long-run equilibrium relationship between the model's variables; $\gamma_{E,i}$ and $\gamma_{f,i}$ correspond to short-term dynamics; the ρ_i coefficient indicates the speed of adjustment of $\tilde{p}_{it}$ to the equilibrium, and its statistical significance indicates a cointegrating relation. If $\rho_i = 0$, there is no cointegration in the model, and Equation (9) is a first-differenced model. Otherwise, there is a return to equilibrium over time after a shock. γ_i^E and β_i measure the short- and long-run PTM, respectively.

Also we can express the bilateral exchange rate between the currencies of the destination market (F) and that of the exporting country (D) through the triangular relation among two exchange rates:

$$E_i = F_{USD} \times USD_{D} = E_{F,USD} \times E_{USD,D} \tag{10}$$

where *F$* is the foreign currency in the destination market, *USD$* corresponds to the US dollar, and *D$* is the currency of the exporting country, in this case, the real (*R$*). We can test the PTM hypothesis by taking into account the possibility that exporters use dollar to invoice exports. Rewriting the long-term relationship (5) as:

$$\tilde{p}_{it} = \alpha_i + \lambda_i' f_t + \beta_{1,i}\tilde{E}_{F,USD,t} + \beta_{2,i}\tilde{E}_{USD,D,t} + u_{it} \tag{11}$$

If firms set their export prices in dollars considering prevailing prices in the international market as a reference, this would be equivalent to the null hypothesis of markup adjustments to follow international prices in dollars regardless of the destination market. That is, considering export prices in domestic currency (R$), then we should expect $\beta_{1,i} = 0$ and $\beta_{2,i} = -1$ for all i in Equation (11) under the null hypothesis. That is, when prices are fixed in dollars, the markup in domestic currency must seek a complete adjustment to the exchange rate variations of the dollar in real units and must respond only to cost variations. We are not able to test this hypothesis for each destination market i individually, but we can test the joint null hypothesis that $E(\beta_1) = 0$ and $E(\beta_2) = -1$ for each equation by product through a joint restriction test [9]. This hypothesis is based on the evidence from Gopinath *et al.* (2020) and Boz *et al.* (2022) that the bilateral - importer vs exporter - exchange rate matters less than the exchange rate of the currency of the exporting country against the dollar for exchange rate pass-through. The corresponding specification for the ECM considering the presence of unobservable common factors is given by:

$$\Delta\tilde{p}_{it} = \alpha_i + \rho_i(\tilde{p}_{it-1} - \beta_{1,i}\tilde{E}_{F,USD,t-1} - \beta_{2,i}\tilde{E}_{USD,D,t-1} - \lambda_i f_{t-1}) + \gamma_{i,1}\Delta\tilde{E}_{F,USD,t} + \gamma_{i,2}\Delta\tilde{E}_{USD,D,t} + \gamma_{f,i}\Delta f_t + \varepsilon_{it} \tag{12}$$

where $\beta_{1,i}$ and $\beta_{2,i}$ represent long-run PTM coefficients, meanwhile $\gamma_{1,i}$ and $\gamma_{2,i}$ are the short-run PTM coefficients. This analysis considers the triangular relationship to estimate PTM and is our contribution.

4. Econometric methodology

The combination of heterogeneity of the response, variable nonstationarity, and CD can lead to severe distortions in standard panel estimators (Andrews, 2005; Phillips & Sul, 2003, 2007; Sarafidis & Wansbeek, 2012). Additionally, Eberhardt and Bond (2013) and Eberhardt and Teal (2019) present evidence that the fixed-effect, pooled and MG estimators present bias in finite samples and a loss of precision in the presence of an unobserved common factors with heterogeneous factor loadings. To estimate the ECM, we use estimators considering CD based on the multifactor structure, which are based on the CCE estimator proposed by Pesaran (2006). In the CCE estimator, the cross-section averages of the dependent and independent variables approximate the linear combination of the common factors, such that we represent the regression model as:

$$\Delta\tilde{p}_{it} = \alpha_i + \rho_i(\tilde{p}_{it-1} - \beta_i\tilde{E}_{it-1} - \lambda_{1,i}\bar{\tilde{p}}_{it-1} - \lambda_{2,i}\bar{\tilde{E}}_{it-1}) + \gamma_{E,i}\Delta\tilde{E}_{it} + \gamma_{f1,i}\Delta\bar{\tilde{p}}_{it} + \gamma_{f2,i}\Delta\bar{\tilde{E}}_{it} + \varepsilon_{it} \tag{13}$$

We consider two estimators: (i) the CCEP and (ii) the CCEMG. Equation (13) is based on CCEMG estimator, given that the coefficients are indexed by i. The CCEMG estimator calculates the average of the coefficients estimated by the regression above for each unit i of the panel. Under the assumption of unrestricted slope and error variance heterogeneity, we compute the CCEMG estimates as simple averages of country-specific estimates from product-specific regressions and we make inferences about $E(\beta_i) = \beta$ [10]. According to Chudik, Pesaran, and Tosetti (2011), those estimators do not require knowledge of the weak CD of the error term – provided it is sufficiently weak - nor knowledge of the serial correlation of the error term [11].

Pesaran (2015) recalls that the CCE estimator was not originally formulated to deal with lagged dependent variables or weakly exogenous regressors such as in the dynamic model. Chudik and Pesaran (2015b) demonstrate that the presence of lagged dependent variables produces bias in finite samples with panel data and the approximation of common factors by the average of the variables should not be used, which lead to infinite lag-order relationships between unobserved common factors and cross-section averages of the observables when N is large. To correct the bias problem in finite samples, the authors propose the addition of enough lags of cross-section averages in individual equations of the panel. The number of lags of cross-section averages must be at least as large as the number of unobserved common factors. In practice, we do not know the number of unobserved factors. They propose a rule of thumb including a number of lags equal to $p_T = \text{int}\left(T^{\frac{1}{3}}\right)$ for the cross-section averages, in which *int* is the integer and T is the time series dimension. Doing this, they demonstrate that CCEMG performs adequately estimating a dynamic model with heterogeneous coefficients. Thus, we use the CCEMG estimator with the addition of lags of cross-section averages to the ECM as in Equation (9), similar to Eberhardt and Presbitero (2015).

In order to compare with the estimates provided by the CCE estimators, we initially present the estimation of models without correction for common factors. So, we also have the following estimators: two-way fixed effects (2FE), Pooled Mean Group (PMG) – see Pesaran, Shin, and Smith (1999) – and Mean Group (MG) – see Pesaran and Smith (1995). The PMG estimator does not allow for the heterogeneity of long-term coefficients, while the MG allows for both the short and long-term. Knetter (1989) considers time dummy variables to capture changes in the marginal cost that would be homogeneous in different destination markets and we include such variables to estimate by 2FE. However, the PMG and MG estimators are not compatible with the adoption of the identification strategy in Knetter (1989), given that these estimators do not allow the inclusion of time dummy variables to capture the homogeneous common factor. Then, we use all variables in deviation from the cross-section average to estimate by PMG and MG controlling for the common temporal effect of Knetter (1989)'s identification strategy.

The estimation of dynamic models will be accompanied by the CD test from Pesaran (2021), the exponent CD from Bailey, Kapetanios, and Pesaran (2016) and the CIPS unit root test from Pesaran (2007) for the residuals to determine the statistical adequacy of the models. Pesaran (2021) test for CD evaluates the presence of CD in the model residuals, while the exponent CD analyzes the strength of CD. Bailey *et al.* (2016) consider that $\Psi \geq 0.5$ indicates strong CD, in which Ψ is a constant that is the exponent of CD or the strength of the unobserved common factors.

5. Database and descriptive statistics

PTM estimation studies often observe sample restrictions due to the nature of export unit value data. Due to variations in data classification or excessive presence of missings or outliers, studies of PTM have relatively small samples [12]. As we have a dynamic model that takes into account the long-term relationship between the variables and due to the asymptotic properties of the CCE estimator, we aim to follow two basic criteria for the construction of the database. The first criterion is a significant volume of exports to each destination over the entire sample. The second criterion is to choose products with the largest number of destination markets considering the entire sample period to approximate the dimensions of countries and time in a balanced sample so that it improves the convergence properties for the heterogeneous estimator [13]. Applying these criteria and considering the quality of available information, our sample contains 25 product categories and 18 destination countries. The period covers the first quarter of 2010 to the fourth quarter of 2020 in a total of 44 quarters.

We list the products and their respective codes in Table 1. The products in the sample are differentiated according to the classification of Rauch (1999), which are the goods that have the lowest probability of predominance of the coalescing effect for exports according to Goldberg and Tille (2008) [14]. These 25 products represent 2.9% of Brazilian exports. Exports to the 18 selected countries account for between 38.4% and 85.6% of the total exported for the selected products, with an average share equal to 60.1% of the total exported.

Code	Product	% total exports
392690	Articles of plastics and articles of other materials of heading	65.09
401693	Gaskets, washers and other seals, of vulcanised rubber	56.30
401699	Articles of vulcanised rubber	77.03
640299	Footwear with outer soles and uppers of rubber or plastics	56.74
640399	Footwear with outer soles of rubber, plastics or composition leather, with uppers of leather	44.65
731815	Threaded screws and bolts, of iron or steel, whether or not with their nuts and washers	66.53
731816	Nuts of iron or steel	64.35
731822	Washers of iron or steel	63.90
731829	Non-threaded articles, of iron or steel	44.82
732020	Helical springs, of iron or steel	77.85
840991	Parts suitable for use solely (with spark-ignition internal combustion piston engine)	54.52
840999	Parts suitable for use solely (with compression-ignition internal combustion piston engine)	54.97
841330	Fuel, lubricating or cooling medium pumps for internal combustion piston engine	70,00
841430	Compressors for refrigerating equipment	58.54
842199	Parts of machinery and apparatus for filtering or purifying liquids or gases	76.16
848210	Ball bearings	69.41
848310	Transmission shafts, incl. cam shafts and crank shafts and cranks	38.45
848330	Bearing housings for machinery, not incorporating ball or roller bearings	40.32
850152	AC motors, multi-phase, of an output	41.28
853690	Electrical apparatus for switching electrical circuits	51.58
854442	Electric conductors for a voltage, insulated, fitted with connectors	46.26
870829	Parts and accessories of bodies for tractors, motor vehicles	85.56
870830	Brakes and servo-brakes and their parts, for tractors, motor vehicles	72.57
870899	Parts and accessories, for tractors, motor vehicles for tractors, motor vehicles	66.03
903289	Regulating or controlling instruments and apparatus	59.26

Note(s): % Total exports reports the ratio between the value exported by Brazil to the 18 countries in the sample and the total exports of that product by Brazil
Source(s): Table by authors

Table 1. Products in the sample and the importance of the countries analyzed for Brazilian exports of that product

So, we have as destination markets: Argentina, Bolivia, Chile, China, Colombia, Costa Rica, Dominican Republic, France, Germany, Guatemala, Italy, Japan, Mexico, Paraguay, Peru, South Africa, UK and Uruguay [15]. We calculate the quarterly average of the monthly series to obtain the series at a quarterly frequency.

This study uses the unit value that is the ratio between the export value and the export quantity. The logarithm of the unit value is a proxy of the logarithm of the export price $\tilde{p}_{it}$. The United Nations International Trade Statistics Database (UN-COMTRADE) makes available the export value and quantity series up to six digits with the Harmonized System. Based on Knetter (1989), we convert export values to the domestic currency (Brazilian real) using the monthly average of the nominal exchange rate.

The UN-COMTRADE compiles the exported value and exported quantity data based on information reported by each member country. This decentralized process of supplying data sometimes results in inconsistent values, generating outliers in the series. To mitigate the effect of this type of outlier on the estimation, we treat all the unit value series using an iterative outlier detection and adjustment procedure from Chen and Liu (1993). We only consider additive-effect (AO) outliers, characterized by isolated peaks that may be due to a measurement error in the series.

One of the regressors in the model is the logarithm of the bilateral nominal exchange rate of the currency of each importing country i in relation to the exporting country's currency – the real (R$) – using data from the International Monetary Fund's International Financial Statistics. In addition, we use the consumer price index series of the destination market to deflate the bilateral exchange rate. We extract these data from the Organization for Economic Cooperation and Development's Statistics. We also use the logarithm of the bilateral exchange rate series of the currency of the destination country against the dollar $\tilde{E}_{F,USD,t}$ from the same data source. Finally, we utilize the logarithm of the bilateral exchange rate of the dollar against the exporting country's currency $\tilde{E}_{USD,D,t}$ from the Brazilian Central Bank.

Table 2 presents the descriptive statistics of the dependent variable $\tilde{p}_{it}$ and the explanatory variables $\tilde{E}_{it}$, $\tilde{E}_{F,USD,t}$ and $\tilde{E}_{USD,D,t}$ of Equations (9) and (12). Overall standard deviation is the common measure of the standard deviation of a sample and is the standard deviation for all the panel data without considering the individual or time dimension. On the one hand, within standard deviation presents the variation of the variable within each individual, while it ignores all variations between individuals. On the other hand, between standard deviation presents the variation of the variable between individuals, while it ignores

Variable		Average	Std. Dev	Min	Max
$\tilde{p}_{it}$	overall	3.933	1.127	−0.466	7.438
	between		0.834	2.723	5.985
	within		0.731	−1.173	6.264
$\tilde{E}_{it}$	overall	0.771	2.703	−5.544	4.663
	between		2.749	−4.749	4.057
	within		0.403	−0.252	1.610
$\tilde{E}_{F,USD,t}$	overall	1.796	2.675	−3.867	5.226
	between		2.749	−3.725	5.082
	within		0.105	1.430	2.077
$\tilde{E}_{USD,D,t}$	overall	−1.025	0.357	−1.686	−0.467
	between		0.000	−1.025	−1.025
	within		0.357	−1.686	−0.467

Source(s): Table by authors

Table 2.
Descriptive statistics of the panel data

all variations over time of each individual (Cameron & Trivedi, 2010). The overall variation of the logarithm of the unit value is equal to 1.13, which is broken down almost equally in between and within variations. The overall variation of the logarithm of the bilateral exchange rate, $\tilde{E}_{it}$, is predominated by the between variation and the same is true for bilateral rates against the dollar ($\tilde{E}_{F,USD,t}$). This means that there is more variation across countries than over time for these variables. Based on the within variation, most of the variation over time of $\tilde{E}_{it}$ is from the exchange rate between the dollar and the real ($\tilde{E}_{USD,D,t}$). That is, the main source of variation in the bilateral exchange rate between the currencies of the destination market and the exporting country over time comes from the variation in the exchange rate value of the dollar against the currency of the exporting country.

Table 3 reports the correlations between the variables of the logarithm of the price ($\tilde{p}_{it}$), the exchange rate between the currencies of the destination and export market ($\tilde{E}_{it}$), of the destination market in US currency units ($\tilde{E}_{F,USD,t}$) and the US currency in Brazilian real ($\tilde{E}_{USD,D,t}$). The correlation between the logarithm of the price and the logarithm of the exchange rate of the dollar in the Brazilian real is −0.63 for the level of variables, with the strongest relationship between price and exchange rate. On the other hand, the correlation is approximately zero when we consider that the logarithm of the price has a correlation of −0.01 to the logarithm of the bilateral exchange rate between the currencies of the destination and export markets and a correlation of 0.07 to the logarithm of the bilateral rate of the currency of the destination market in terms of the dollar. The correlations between the variables in the first difference show that most of the association between the change in the logarithm of prices and the change in the bilateral rate between the currencies of the destination and the exporter (−0.23) is due to the association between the change in the logarithm of prices and the change in the logarithm of the exchange rate of the dollar expressed in Brazilian real (−0.27). The highest correlation is between the change in the logarithm of the bilateral exchange rate between the currencies of the destination and export market and the change in the logarithm of the exchange rate of the dollar in Brazilian real (0.92).

The negative correlations between the variables of the logarithm of the price ($\tilde{p}_{it}$) and the exchange rate of the dollar in Brazilian real – both in level and in difference – indicate a movement of compensation from the price in Brazilian real to changes in the dollar against Brazilian real to stabilize prices in dollars over time. Figure 1 shows the average of the logarithm of the unit value of exports by product in dollars (on the left) and in R$ (on the right). The prices of products in dollars fluctuate around a constant over time on the left side of this figure, while the prices of exports of products in R$ show an upward trend over time on the right side. In other words, there seems to be a movement of markup compensation associated with export prices that are consistent with the predictions of price stabilization in the dollar.

	Level			First difference		
	$\tilde{p}_{it}$	$\tilde{E}_{it}$	$\tilde{E}_{F,USD,t}$	$\tilde{p}_{it}$	$\tilde{E}_{it}$	$\tilde{E}_{F,USD,t}$
$\tilde{E}_{it}$	−0.01			−0.23		
$\tilde{E}_{F,USD,t}$	0.07	0.99		−0.08	0.75	
$\tilde{E}_{USD,D,t}$	−0.63	0.15	0.01	−0.27	0.92	0.43

Source(s): Table by authors

Table 3. Correlation between the variables in level (on the left) and in first difference (on the right) considering the average price for the country

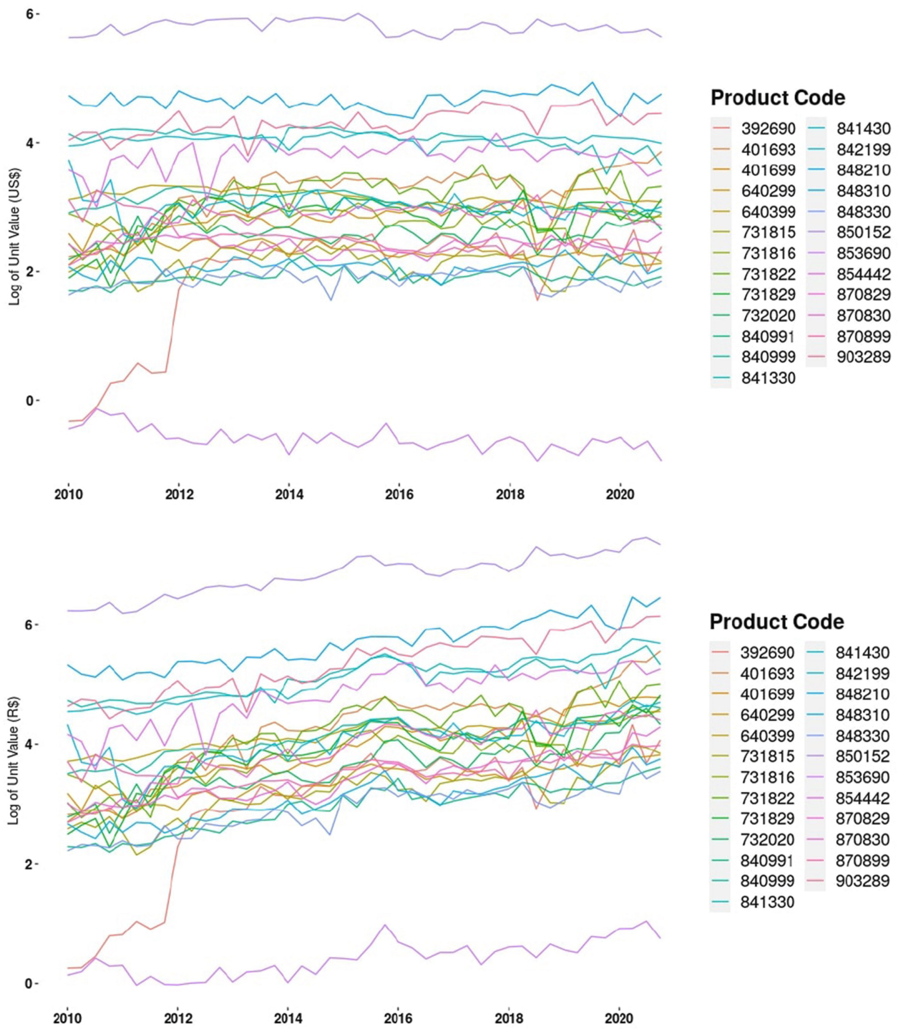

Figure 1.
Average of unit value logarithm in dollars (on the left) and in R$ (on the right) per product over time

6. Results

This section presents the long- and short-run estimates of PTM by the 2FE, PMG, MG, CCEP and CCEMG estimators. Before estimating the ECM, we conduct Gengenbach, Urbain, and Westerlund (2008) test for cointegration between the model variables and we present these results in Table A1 of Appendix. For all the products, we reject the null hypothesis of no cointegration between the variables, leading us to conclude that there must be a representation in the form of an ECM expressed by Equations (9) and (12). Equation (9) has a bilateral exchange rate between the destination market and the exporting country as a regressor. Equation (12) includes the exchange rate of the currency of the destination market relative to the dollar and the exchange rate between the dollar and the currency of the exporting country as regressors. The results of the tests for cointegration also denote the existence of a mean reversion between prices and the nominal exchange rate, indicating that a markup adjustment must take place to ensure the stationarity of the export prices in foreign currency.

We estimate Equations (9) and (12) for each of the 25 products. As proposed by Chudik and Pesaran (2015b), we add three lagged cross-section averages to correct the small sample bias – $p_T = \text{int}\left(44^{\frac{1}{3}}\right) = 3$, where 44 is the number of quarters. We present the estimates of Equations (9) and (12), respectively, in Tables 4 and 5. Table A2 in Appendix reports the individual estimates of the long-term coefficients of Equations (9) and (12).

Table 4 summarizes the results of the estimates of Equation (9) by product. The top of Table 4 reports the results of the *CD* tests for cross-section dependence, the exponent CD and the CIPS unit root tests for the residuals of the models. We reject the hypothesis of the presence of a unit root for the residuals of the model for all 25 products using all the estimators at a 10% statistical significance level based on the unit root test. However, the *CD* tests reveal a large discrepancy between the estimators in relation to the cross-section independence of the residuals. We verify the greatest evidence of the cross-section independence of the residuals in the models estimated using the CCEMG estimator at a 10% statistical significance level – for 14 products. Also, we consider the strength of unobserved common factors from the estimate for Ψ, where $\widehat{\Psi} < 0.5$ represents semiweak CD or cross-section independence of the residuals. In general, CCEP and CCEMG have greater cross-section independence. These tests show that the inclusion of common factors and the heterogeneity of the coefficients are jointly important to guarantee the statistical adequacy of the models. This result is similar to that obtained by Eberhardt and Presbitero (2015). Thus,

	2FE	PMG	MG	CCEP	CCEMG
			Diagnosis of the residuals		
Independent residuals (CD)	4	8	11	9	14
Independent residuals (Exponent)	13	8	7	17	17
Stationary residuals	25	25	25	25	25
			Speed of adjustment (ρ_i)		
Average	−0.66	−0.27	−0.75	−0.18	−0.75
Median	−0.64	−0.21	−0.75	−0.16	−0.76
Standard deviation	0.19	0.18	0.11	0.10	0.12
Maximum	−0.30	−0.05	−0.52	−0.05	−0.50
Minimum	−1.00	−0.83	−0.97	−0.44	−0.96
Significant coefficients at 5%	25	1	25	0	25
			Long-run estimates of PTM (β_i)		
Average	−0.17	0.03	−0.13	0.00	−0.91
Median	−0.23	0.02	−0.21	0.01	−0.93
Standard deviation	0.65	0.08	1.11	0.08	0.27
Maximum	1.15	0.22	2.63	0.18	−0.33
Minimum	−1.19	−0.09	−1.79	−0.13	−1.52
Significant coefficients at 5%	4	0	1	0	24
Reject H_0 ($\beta_i = -1$) at 5%	17	17	9	20	6
			Short-run estimates of PTM (γ_i^E)		
Average	0.26	−0.05	−0.13	0.23	0.58
Median	0.14	−0.07	−0.21	0.33	0.50
Standard deviation	0.55	0.79	1.11	0.70	1.00
Maximum	1.61	1.57	2.63	1.66	2.85
Minimum	−0.60	−1.57	−1.79	−0.95	−0.75
Significant coefficients at 5%	0	2	2	0	3

Source(s): Table by authors

Table 4. Summary of the estimates by the different estimators and the diagnosis of the residuals for the 25 products using the deflated exchange rate as regressor

Table 5. Summary of the estimates by the different estimators and the diagnosis of the residuals for the 25 products using 2 exchange rates as regressors

	2FE	CCEP	CCEMG
		Diagnosis of the residuals	
Independent residuals (CD)	4	4	16
Independent residuals (Exponent)	12	23	20
Stationary residuals	25	25	25
		Speed of adjustment (ρ_i)	
Average	−0.66	−0.24	−0.84
Median	−0.64	−0.20	−0.85
Standard deviation	0.19	0.14	0.10
Maximum	−0.30	−0.06	−0.59
Minimum	−1.00	−0.63	−1.00
Significant coefficients at 5%	25	0	25
		Long-run estimates of PTM ($\beta_{1,i}$)	
Average	−0.17	0.02	0.47
Median	−0.23	0.01	0.41
Standard deviation	0.65	0.08	1.22
Maximum	1.15	0.21	4.99
Minimum	−1.19	−0.10	−1.19
Significant coefficients at 5%	4	0	2
		Long-run estimates of PTM ($\beta_{2,i}$)	
Average	−1.16	−1.10	−1.05
Median	−1.10	−1.02	−1.06
Standard deviation	0.45	0.41	0.30
Maximum	−0.58	−0.43	−0.52
Minimum	−2.17	−2.45	−1.67
Significant coefficients at 5%	17	6	25
Reject H_0 ($\beta_{2,i} = -1$ and $\beta_{1,i} = 0$) at 5%	4	0	4
		Short-run estimates of PTM ($\gamma_{1,i}$)	
Average	0.27	0.23	0.10
Median	0.15	0.21	−0.27
Standard deviation	0.55	0.62	1.19
Maximum	1.61	1.51	3.72
Minimum	−0.60	−0.57	−1.98
Significant coefficients at 5%	0	0	1
		Short-run estimates of PTM ($\gamma_{2,i}$)	
Average	−7.15	0.03	0.66
Median	−3.69	0.00	0.64
Standard deviation	41.96	0.15	0.37
Maximum	60.97	0.45	1.80
Minimum	−140.29	−0.25	0.02
Significant coefficients at 5%	2	0	5

Source(s): Table by authors

we can state that the CCEMG estimator provides the most appropriate estimate for the dynamic model.

Concerning to the speed of adjustment estimates, we have evidence of cointegration through the statistical significance of the error correction term, ρ_i. The speed of adjustment coefficients are statistically significant at 5% for all products in the case of estimators that allow heterogeneous coefficients – MG and CCEMG – and the 2FE estimator. The absolute value of the average error correction coefficients is 0.66 for the 2FE estimator and 0.75 for the CCEMG estimator. The absolute value of 0.75 for the coefficient of the speed of adjustment indicates that the time needed in order to eliminate 50% of the deviation in the long-term relationship is just 0.5 quarter (half-life is 0.5) [16]. The heterogeneity of the coefficients obtained by the CCEMG estimator leads to the highest speed of adjustment. The high speed of adjustment may be related to the non-statistical significance of the short-run PTM coefficient, coupled with the possibility of price stickiness in the short run. Our result is in line with Gagnon and Knetter (1995), who obtained that PTM is greater in the long run than in the short run for the automobile industry, with the exception of exports to the USA and Canada. Pooled estimators – PMG and CCEP – do not indicate cointegration or the practice of PTM. The results of the pooled estimators are divergent and the models do not present independent residuals in cross-section in most cases according to the CD tests.

We also present a summary of the estimates of the long-run PTM coefficient, $\widehat{\beta}_i$, in Table 4. Table A2 in Appendix reports the long-term PTM estimates $\widehat{\beta}_i$ for the different estimators by product. There is a large divergence in estimates. While the 2FE, MG and CCEMG estimators present negative coefficients, favoring the LCPS hypothesis, the long-term PTM estimates are statistically significant at 5% only for 4 products when produced by the 2FE estimator, for 1 product by the MG estimator and for 24 products by the CCEMG estimator. On the other hand, the PMG and MG estimators present coefficients with values close to zero and not statistically significant at 5%.

The statistical significance of the PTM coefficient for most products distinguishes the estimates by CCEMG from the other estimators. The average of the estimates produced by the CCEMG estimator is -0.91 for $\widehat{\beta}_i$. In other words, a 10% depreciation of the exporter's currency relative to a particular market leads to a 9.1% destination-specific reduction in the markup of the export price over marginal cost. Destination-specific adjustments in markup smooth the effect of a bilateral exchange rate change on the price in units of the importer's currency. Furthermore, we reject the null hypothesis that the long-term PTM coefficient is equal to -1 in only 6 products at a statistical significance level of 5% by the CCEMG estimator. In other words, the markup in domestic currency does not adjust completely to cancel out the price variation produced by the exchange rate change only in the case of six products.

The summary of the short-run estimates of PTM using the different estimators is in Table 4. The CCEMG estimator produces coefficients statistically significant at 5% only for three products. In most of the estimates produced by all estimators, most product prices do not respond to exchange rate fluctuations in the short term. This evidence seems consistent across estimators.

In order to test the PTM hypothesis considering the triangular relation of the exchange rate and the vehicle currency (USD), we estimate Equation (12). Table 5 reports the estimates by decomposing the exchange rate. We do not use the PMG and MG estimators to estimate Equation (12) because these estimators do not allow including the regressor $\tilde{E}_{USD,D,t}$ (which is invariant in the cross-section dimension) and controlling for time effects – we cannot use all variables in deviation from the cross-section mean with this specification. Regarding the quality of the residuals, the results in this second round of estimation are even better than those of the first round with 16 of the 25 products estimated by the CCEMG estimator with independent cross-section residuals by CD test. The CCEMG estimator continues to offer a better alternative for the estimation in terms of the statistical adequacy of the model by CD test. The CCEP estimator leads by the exponent cross-section dependence, but with a small difference in relation to CCEMG estimator. Considering the set of statistical adequacy tests (CD test and exponent CD), the most appropriate results are those by the CCEMG estimator.

The estimates of the error correction speed coefficient are a source of discrepancy between the CCEP and CCEMG estimators. The estimates produced by the 2FE and CCEMG estimators are statistically significant at 5% in all products, while the estimates by the CCEP estimator do not indicate the presence of cointegration in any of the cases. The

Gengenbach *et al.* (2008) test for cointegration presented in Table A1 of Appendix indicates a different result in relation to the CCEP estimator.

The CCEMG estimator produces an average for the error correction speed of adjustment coefficient, $\widehat{\rho}_i$, of -0.84 and it is statistically significant at 5% for all products. This absolute value is greater than that estimated using the exchange rate between the currencies of the destination market and the exporting country (-0.75). Thus, the speed of adjustment toward equilibrium is greater when decomposing the exchange rate. The time needed to eliminate 50% of the deviation in the long-term relationship is just 0.38 quarter.

When we decompose the exchange rate, the coefficient associated with the exchange rate between dollar and real $(\tilde{E}_{USD,D,t})$ is statistically significant at 5% for all products with the CCEMG estimator and with an average of -1.05 among the products. The average coefficient estimated by the CCEMG is not different compared to those from the 2FE and CCEP estimators, but these estimators have a smaller number of products with statistically significant at 5% [17]. We note that the average of long-run estimates of PTM by 2FE (Table 4) is equal to -0.17, while this average $(\overline{\beta}_{2,i})$ is -1.16 when decomposing the exchange rate (Table 5). The estimation method is not the only difference in the result, but this change is mainly due to the specification adopted.

In the estimation by CCEMG, the coefficient associated with the exchange rate of the dollar against the domestic currency is statistically significant at 5% in all products, while by 2FE in 17 of the 25 products considered. We do not reject the joint null hypothesis $\beta_{1,i} = 0$ and $\beta_{2,i} = -1$ at a significance level of 5% in 21 of the 25 cases. Using as a covariate the exchange rate of the currency of the destination market in relation to the Brazilian real, we do not reject the null hypothesis that $\beta_i = -1$ at the same level of significance for 19 products by CCEMG. The greatest difference on the same hypothesis appears in the results of the 2FE estimator. Previously we rejected the null hypothesis in 17 cases, while we reject this null hypothesis in only 4 products using the exchange rate of dollar per Brazilian real. That is, despite the difference in the estimator used, the evidence indicates that what matters for the long-term pricing of the Brazilian exporter is the exchange rate of dollar per Brazilian real. This evidence is in line with the exchange rate pass-through literature based on Gopinath *et al.* (2020) and Boz *et al.* (2022) that bilateral – importer vs exporter – exchange rate matters less than the exchange rate of the exporting country's currency against the dollar for exchange rate pass-through.

As before, the short-term coefficients for both exchange rates ($\tilde{E}_{F,USD,it}$ and $\tilde{E}_{USD,D,it}$) are not statistically significant at 5% by the different estimators for most products. This result is robust regardless of the specification used. But the coefficient associated with the exchange rate of the dollar against the Brazilian real in the short term has statistical significance at 5% in a greater number of products than that associated with the exchange rate between the currencies of the importer and the exporter – five against three cases previously.

6.1 Heterogeneity of the PTM coefficients

In this section, we study the patterns regarding heterogeneity of the coefficient estimates for the long-term PTM produced by the CCEMG estimator. Tables 1 and 2 in the Supplementary Appendix report, respectively, the individual estimates of β_i and $\beta_{2,i}$ of Equations (9) and (12), that we analyze in this subsection [18]. We present the descriptive statistics of these two panels – in which the dimensions are by product and by country – in Table 6.

Initially, we observe in Table 6 that although the averages of the coefficients are very close (-0.91 and -1.05), there is a distinct pattern in the other statistics. When we look at the standard deviation statistic and its decomposition, the coefficients $\widehat{\beta}_i$ estimated with Equation (9) present a greater variability in both dimensions. The bottom of Table 6 shows

Variable		Mean	Std. Dev	Min	Max
$\widehat{\beta}_i$	overall	−0.91	1.24	−6.99	18.14
	country		0.31	−1.71	−0.49
	product		1.20	−7.20	17.92
$\widehat{\beta}_{2,i}$	overall	−1.05	1.10	−8.91	3.24
	country		0.22	−1.57	−0.73
	product		1.08	−8.38	3.17

Correlation between average PTM coefficients by country and by product

	$\overline{\beta}_{i,country}$	$\overline{\beta}_{i,product}$
$\overline{\beta}_{2,i,country}$	−0.33	
$\overline{\beta}_{2,i,product}$		0.79

Source(s): Table by authors

Table 6. Descriptive statistics of the PTM coefficients in the country and product dimensions as a panel and correlation between PTM coefficients

the correlation coefficients between the averages of $\widehat{\beta}_i$ and $\widehat{\beta}_{2,i}$ in the country and product dimensions. The difference between the two estimates ($\widehat{\beta}_i$ and $\widehat{\beta}_{2,i}$) comes from the country dimension due to the negative correlation between the two coefficients (−0.33).

Figure 2 illustrates the averages of the coefficients calculated in the product and country dimensions for the two estimated models. Graphically, the specification with the triangular decomposition of the exchange rate has different results in the country dimension than that with the exchange rate between the currencies of the importer and the exporter. While the results by-products are similar on average, the specification using two exchange rate regressors tends to produce greater homogeneity of coefficients in the country dimension, bringing them closer to −1 on average. This greater homogeneity in the country dimension with the triangular decomposition of the exchange rate must be related to the point that firms set export prices in the dollar considering global demand rather than specific destination markets as Corsetti *et al.* (2018) and Gopinath *et al.* (2020).

We obtain a complementary result in Table A3 of Appendix, which presents the estimates of two equations in which the dependent variables are, respectively, $\widehat{\beta}_i$ and $\widehat{\beta}_{2,i}$. We have as regressors the logarithm of Gross Domestic Product (GDP) per capita and dummy variables of product and country [19]. To save degrees of freedom, we use the product dummy variables at the 4-digit level to control for heterogeneity in the product dimension rather than at the 6-digit level. In the first result for $\widehat{\beta}_i$, we observe that in addition to the statistical significance of the GDP per capita variable, three product dummy variables and nine country dummy variables are statistically significant at 5%. In the equation where the dependent variable is $\widehat{\beta}_{2,i}$, we only observe two statistically significant coefficients for product dummy variables and two coefficients associated with country dummy variables at 5%. This is further evidence for the greater homogeneity of the export price between the destination countries when considering the exchange rate of the dollar against the real instead of the exchange rate between the currencies of the destination and exporting countries. This result indicates that firms set the price of the product in dollar in our case.

7. Discussion of results

Initially, we have to remember that a large part of the Brazilian export basket consists of basic and semi-manufactured products, which are strongly influenced by supply and demand on

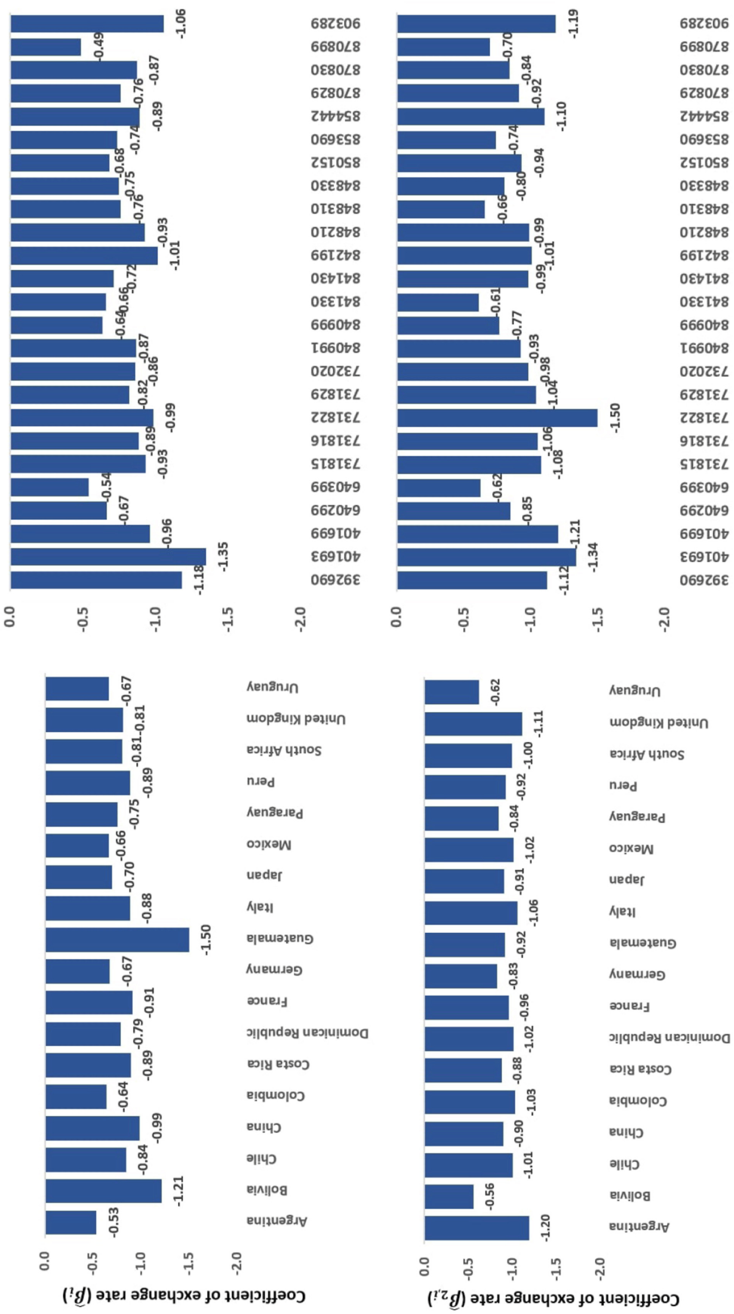

Figure 2.
PTM coefficients by country (on the left) and by product (on the right) using real exchange rate (at the top) and nominal exchange rate between US dollar and exporting country currencies (at the bottom) as regressors

international markets [20]. Hence, the capacity of domestic exporters to determine prices in the global market is diminished, but this aspect is not the central focus of our analysis. As a result, we have opted for a selection of manufactured products that encompass a diverse range of export destinations. This allows us to observe both sources of variation of PTM estimates, the variation of product and destination market. The 25 selected products, as previously mentioned, represent an average of 2.9% of total Brazilian exports for the period, representing six sectors of the manufacturing industry [21]. Nevertheless, we acknowledge that the results obtained here cannot be extrapolated to the entire range of industrialized products in the export basket.

PTM estimates indicate that Brazilian exporters set their products' prices in US dollars on the international market for 24 of 25 products. That is, unlike the "Producer Currency Pricing" or "Local Currency Pricing" hypotheses, the analyzed exports would be best characterized as vehicle currency price setters, that is, "Vehicle Currency Pricing." By integrating theory and additional empirical evidence on exchange rate pass-through, we can argue the outcomes and implications.

Considering a set of stylized facts present in the international and national literature about exporting firms, we can also relate our results to export pricing strategies [22]. Based on a cross-section of countries, including Brazil, these facts are that exporting firms:

(1) are a minority in the set of firms, which export only a small fraction of their production;

(2) tend to be more productive and larger, but heterogeneous, with a positive relationship between size, productivity differential and number of destination markets;

(3) the distribution of exported value is asymmetric toward the largest exporting firms, and these firms are also those with the highest import intensity.

From a theoretical point of view, the first observation is the convergence of Brazilian export prices to an international price system. This brings us back to the idea of a "small economy." Under this hypothesis, exporters lose the ability to set prices on international markets, considering these prices as given. However, unlike the small economy hypothesis, the adoption of an international price reference system is an endogenous process, driven by the exporter's choice when determining prices in a vehicle currency. According to Corsetti *et al.* (2018), companies would establish a global price in dollars for their products, maximizing their profits in relation to global demand taken as a whole. Price in dollars overcomes market segmentation and translates into a "Reference Pricing System," whereby companies do not exploit specific market demand elasticities, but price in relation to global demand [23].

The predominance of export prices set in US dollars makes exporters' profitability extremely dependent on variations in the exchange rate of the national currency in relation to the dollar. Thus, instead of domestic currency depreciation leading to a perception of cheaper Brazilian industrial exports, what we actually have is price stability in dollars and an increase in profit margins. Thus, there would be no expected increase in demand for domestic exports, which would explain the relative insensitivity of exports to exchange rate fluctuations [24], [25].

However, also taking into account the recent literature on exchange pass-through, which establishes a relationship between exchange rate pass-through and import intensity, we can make additional considerations that help us in this discussion. Gopinath *et al.* (2020) derive predictions for non-US countries that ERPT to import prices should be high and driven by the dollar exchange rate instead of the bilateral exchange rate. The Brazilian case corroborates these predictions by Choudhri and Hakura (2015), Kannebley Júnior, Godoi, and Prince (2022) and Kannebley *et al.* (2023) [26]. Amiti, Itskhoki, and Konings (2014) present evidence for

Belgian firms that the exchange rate pass-through of exporting firms is inversely correlated with the import intensity of these firms. On the other hand, Chung (2016) argues theoretically and presents evidence that exporters' dependence on imported inputs affects their choice of invoicing currency. Chung (2016) demonstrates that exporters who depend more on foreign currency-denominated inputs are less likely to price in their home currency. Their evidence for UK firms is that VCP is more likely for exporters with a higher share of VCP inputs. This set of evidence therefore indicates that imported input costs when denominated in dollars must have a high degree of exchange rate pass-through and that firms are likely to price in the currency of their imported inputs to mitigate the impact of exchange rate fluctuations on marginal costs.

Brazil is still a typical case of using the USD as a vehicle currency in foreign trade transactions and imports. Therefore, given the composition of our sample which favors more technology-intensive sectors and a wider number of destination markets, it is reasonable to assume that exporting firms are those that are more productive and import more intensively. Therefore, we assume that exchange rate variations tend to have an opposite effect on export profitability due to their impact on costs, which would be more dependent on imported inputs denominated in dollars in the case of the products analyzed in our work. As a result, the higher share of imported inputs denominated in US dollars in the total costs would serve as an incentive for exporting firms to adopt a pricing practice based on the US dollar due to the opposing effects of exchange rate fluctuations on costs.

Consequently, a combination of incentives arising from the existence of strategic complementarities – reflecting competitive pressures in the international market where competitors also tend to adopt VCP – and the desire to minimize exchange rate risks associated with imported input costs leads to a pursuit of profit maximization. This pursuit involves maintaining price stability relative to competitors' prices. Under the assumption of an international price system in place, this would discourage discrimination between destination markets. This would contradict the PTM hypothesis for Brazilian manufacturing exports.

8. Conclusion

This article seeks to implement a new econometric approach to estimate the degree of PTM in the case of an emerging economy that predominantly uses the dollar as its vehicle currency. We also explore the hypothesis of the firm setting the international market price in dollars so that the firm does not take into account the bilateral exchange rate between the currencies of the destination and exporting countries. This is because firms expect their competitors to follow the same rationale.

We estimate the degree of PTM in the short and long term with the ECM in panel data using a sample of 25 manufactured products exported by Brazil to 18 countries between 2010 and 2020. Our results indicate that the identification strategy of Knetter (1989) – based on a homogeneous common factor to represent the evolution of the exporter's costs – is not adequate in the presence of global common factors that can lead to inefficient and inconsistent PTM estimates. Additionally, an alternative specification with the triangular decomposition of the exchange rate inspired by Boz *et al.* (2022) and Gopinath *et al.* (2020) produced more enlightening results on the pricing of exports, which is in line with the predictions of models based on the dominant currency paradigm.

In econometric terms, a possible interpretation of our results is that the use of the bilateral exchange rate between the currencies of the destination and exporting countries as an explanatory variable has the same effect as using a variable with measurement error in the case of a country such as Brazil that considers the bilateral exchange rate of the dollar in terms of the domestic currency in your transactions abroad. As part of the exchange rate variation resulting from the fluctuation of the currency of the destination market in dollars is not incorporated in the pricing of exports, this "noise" – which is correlated with the common factor brought by the dollar being a dominant currency in the international market – biases the PTM estimates leading to specification errors in the estimated model.

Our results indicate heterogenous long-term relations across countries. However, these heterogeneous cointegration relationships do not allow us to infer whether or not the markup responds specifically to the exchange rate variation in each destination market. The adjustment keeps export prices in dollars relatively stable compared to their international competitors. Despite the tendency to align with international competitors who also use dollar pricing, there are cases where firms may deviate from this behavior.

In economic terms, given the characteristics of exporting firms and considering the wide predominance of the US dollar in the invoicing of Brazilian trade transactions, we can infer that the pricing of Brazilian exports in dollars reflects a profit maximization strategy that considers an international price system based on global demand for products. This would eliminate market segmentation, which would be in accordance with the practice of pricing exports without market-specific markup adjustment. This means that even though the markups are variable in domestic currency, the set of exported Brazilian products considered in this study has a reduced probability of specific and variable adjustments between destination markets and, therefore, reduced PTM practice. Our result on average is in agreement with the prediction of Corsetti *et al.* (2018) that the extent of destination-specific markup adjustments is lower using a vehicle currency.

Gopinath (2015) argues that the international price system framework has implications for export competitiveness and trade balance adjustment following exchange rate fluctuations. As most countries use the vehicle currency to price their exports, they are relatively insensitive to exchange rate fluctuations. This would make trade balance adjustments, through relative price effects, more likely to be driven by import adjustments. The results produced in our work are evidence favorable to this hypothesis, but which would deserve further studies that involve estimates of impacts on the extensive and intensive margins of Brazilian exports, in addition to a possible expansion of the set of products analyzed.

There would be several ways to extend this work as a research agenda if we had information at the microdata level from exporting firm. In addition, we could discriminate exports by the currency in which they are invoiced using firm data. We could analyze the difference in export price (if any) between destination markets and exchange rate sensitivity – and use triangular decomposition. The discussion on access to data for research goes beyond the scope of this article. Our major limitation is that we use unit value as an approximation of the exported price. We do not observe the exported price series. Furthermore, the data are aggregated by product, but not at the firm level. This leads to aggregation, although there are studies using unit values with harmonized systems in the literature. Our evidence is limited to unit value information without such discrimination by currency, which does not allow for a better identification of the impacts of exchange rate variations on marginal costs and the sensitivity of export prices considering different currency invoices.

A final observation to make is that, although the Brazilian export basket is much larger, most of the products responsible for Brazil's foreign trade surplus are products that belong to the classes of "organized exchange" or "referenced priced" goods based on Rauch (1999) classification. In these cases, pricing in dollars and using competitors' international prices as a reference is even more likely than differentiated goods. In this sense, although we know the sample of products used in this study is relatively limited, they could reflect with accuracy the pricing practice of Brazilian exporters on international markets even in the case of differentiated products as in our sample.

Notes

1. Goldberg & Tille (2008), Gopinath (2015) and Ilzetzki *et al.* (2019) corroborate this evidence.
2. Other factors are inertia in currency use and the best hedge against cost fluctuations arising from relevant movements in aggregate demand and wages in the source and destination countries.
3. Additionally, 90% of the public and publicly guaranteed debt is denominated in USD according to Ilzetzki *et al.* (2019) and 86.03% according to information from the Central Bank of Brazil in https://www.bcb.gov.br/estabilidadefinanceira/relgestaoreservas.
4. The survey covered 7,002 companies in three economic sectors: manufacturing, services and commerce. Of this total, 1,855 firms were exporters.
5. Fitzgerald & Haller (2014) and Lavoie & Liu (2007) point out restrictions to this data. However, Brazil does not disclose foreign trade data of Brazilian firms due to fiscal secrecy, which does not allow us to reproduce empirical strategies from other studies that use microdata from firms.
6. The triangular decomposition is that, instead of using the exchange rate of the destination market currency in relation to the exporter's currency, we use two exchange rates: the first is the destination market currency in relation to the dollar and the second is the dollar against the exporter's currency.
7. In static models, we disregard issues such as adjustment costs, production bottlenecks and uncertainty. The nature of the exchange rate variation – temporary or permanent – is not relevant to generate PTM.
8. Empirical studies generally make use of nominal exchange rates – units of the destination market currency per unit of exporter's currency – deflated by the wholesale price level in the destination market such as Knetter (1989, 1993), Gagnon & Knetter (1995), among others. According to Knetter (1995), the rationale for deflating is that the optimal price charged by the exporter should be invariant to movements in the nominal exchange rate that correspond to inflation in the destination market.
9. As shown below, the estimator used assumes that βi follows the random coefficient model $\beta i = \beta + \eta i$, $\eta i \sim (0, \Sigma\eta)$, where $i = 1, \ldots, N$.
10. Under slope heterogeneity, the CCEMG approach assumes that β follows the random coefficient model, which implies that the regressors are strictly exogenous.
11. According to Chudik and Pesaran (2015a), weak, strong and semi-strong common factors may be used to represent very general forms of cross-sectional dependence. These authors exemplify a representation of spatial processes as a factor process with an infinite number of weak factors and no idiosyncratic errors. Strong factors can be used to represent the effect of the cross-section units that are "dominant" or pervasive, in the sense that they impact all the other units in the sample and their effect does not vanish as N tends to infinity, where N is the cross-sectional dimension. Semi-strong factors may exist if there is a cross-section unit or an unobserved common factor that affects only a subset of the units and the number of affected units rise more slowly than the total number of units. According to Pesaran (2015), the presence of weak or semi-strong factors in errors does not affect the consistency of conventional panel data estimators, but it does affect inference.
12. For example, Kasa (1992) uses monthly data from seven products with 7-digit HS for the period 1978 to 1987. Gil-Pareja (2000) reports estimations with quarterly data from 1988 to 1996 for 26 products. Knetter (1989) considers data from 16 products between 1978 and 1986 with an annual sample of unit values of unit values at the 7-digit industry level for seven industries.
13. The asymptotic convergence of the dynamic CCE estimator occurs at the rate $\sqrt{N}$. The ratio $\frac{N}{T}$ needs to be constant for conducting inference due to the presence of small time series bias (Ditzen, 2018).
14. Rauch (1999) provides a classification of commodities at the three and four-digit SITC level into one of three categories: "organized exchange" good, "referenced priced" good and all other goods are classified as "differentiated." Goldberg & Tille (2008) apply this classification to investigate associations between product categories and invoicing patterns.

15. We exclude two countries from the sample which are Ecuador and the USA as they are destination markets that use the US dollar as their local currency. The inclusion of these two countries would not allow estimating Equation (12).
16. Following Arsova (2021), we calculate the implied half-life as $\frac{\ln(0.5)}{\ln(1+\bar{\rho}_i)}$, where $\bar{\rho}_i$ is the average error correction coefficient between destination markets.
17. The correlations for the estimated coefficient $\widehat{\beta}_i$ of Equation (9) are equal to -0.1 between 2FE and CCEP estimators and 0.01 between 2FE and CCEMG estimator. In the estimates for the coefficient $\widehat{\beta}_{2,i}$ referring to Equation (12), the correlations are 0.18 between 2FE and CCEP and 0.2 between 2FE and CCEMG. The correlations between the estimates produced by CCEP and CCEMG are respectively equal to -0.4 and 0.34 for $\widehat{\beta}_i$ and $\widehat{\beta}_{2,i}$.
18. For the second model with the triangular decomposition of the exchange rate, we present only the estimates of the coefficient $\widehat{\beta}_{2,i}$ associated with the exchange rate of the dollar per Brazilian real given that the coefficient associated with the other exchange rate variable was not statistically significant at 5% in the most of the products.
19. We extract the GDP per capita variable from the World Bank – in constant 2015 US dollars. As the dependent variable is an estimate, we use the estimated dependent variable regression based on Lewis & Linzer (2005). We estimate by feasible generalized least squares (FGLS) with a known function for weight ωi. From the residuals of the ordinary least squares regression, we obtain the squared residuals for each observation w_i^2. In addition, we calculate a measure for the variance $\widehat{\sigma}^2$. So we use $\omega_i = \frac{1}{\sqrt{w_i^2 + \widehat{\sigma}^2}}$ for the FGLS estimator.
20. According to the Brazilian Central Bank (2019), basic products dominated the export basket in 2018, with a 51% share of the exported value, followed by manufactured products with 35% and semi-manufactured products with 14%.
21. The corresponding manufacturing sectors would be (i) Preparation of leather and manufacture of leather goods, travel goods and footwear, (ii) Manufacture of rubber and plastic products, (iii) Manufacture of electrical machinery, apparatus and material, (iv) Manufacture of machinery and equipment, (v) Manufacture of vehicles cars and trailers and (vi) Manufacture of other transport equipment, except motor vehicles.
22. The literature from the 1990s and early 2000s on productivity and exports is rich in evidence. Wagner (2007) provides a survey of this topic for international literature, while Kannebley Júnior (2011) presents it in a similar way to national literature. Other references to national literature are De Negri (2003), Hidalgo & Mata (2009), Gomes & Ellery (2007) and da Silva Catela & Gonçalves (2013) which discuss topics related to the profile of exporting firms, number of destination markets and import coefficients of these firms.
23. Non-segmentation in the foreign market does not necessarily imply the lack of segmentation between the foreign and domestic markets, especially in Brazil, which is considered relatively a closed economy. In fact, Kannebley & Assahide (2017) have already found – through the estimation of the Marston (1990) model – that in the long run an exchange rate devaluation should promote a gap between foreign and domestic prices when quoted in the same currency. They conclude that there would be sensitive price discrimination between domestic and international markets for manufacturing export products. This indicates that export prices would be strongly influenced by prices practiced on the international market.
24. However, exchange rate fluctuations might affect exporting firms' entry and exit. Therefore, variations in exported quantity are derived from an increase in extensive margins.
25. Evidence from Berman, Martin, & Mayer (2012) shows that high-performing French firms react to a depreciation by increasing significantly their markup and by proportionally increasing less their export volume.

26. These authors find a range of estimates from complete pass-through (Choudhri & Hakura, 2015) to degrees of pass-through to import prices between 73% and 76%, respectively, in Kannebley Jr *et al.* (2023) and Kannebley Júnior *et al.* (2022).

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Appendix

Products	Real exchange rate as regressor Cross-section averages Without	With	Two exchange rates as regressors Cross-section averages Without	With
392690	−3.84***	−2.94***	−3.92***	−3.45***
401693	−4.72***	−3.81***	−4.77***	−4.37***
401699	−5.11***	−4.85***	−6.74***	−6.09***
640299	−6.02***	−4.54***	−6.47***	−5.27***
640399	−5.66***	−4.09***	−5.83***	−5.12***
731815	−4.90***	−4.27***	−5.03***	−4.80***
731816	−5.91***	−5.46***	−5.87***	−5.62***
731822	−5.26***	−4.80***	−5.32***	−4.96***
731829	−5.68***	−5.19***	−5.82***	−5.50***
732020	−6.13***	−7.62***	−10.00***	−9.50***
840991	−4.59***	−4.03***	−4.69***	−4.28***
840999	−4.43***	−4.24***	−4.99***	−4.77***
841330	−5.23***	−4.55***	−5.77***	−5.15***
841430	−4.37***	−3.99***	−5.19***	−4.97***
842199	−6.09***	−5.77***	−6.24***	−6.05***
848210	−5.39***	−4.82***	−5.34***	−4.96***
848310	−5.11***	−5.13***	−5.70***	−5.30***
848330	−5.25***	−5.23***	−6.13***	−5.76***
850152	−5.76***	−4.83***	−5.91***	−5.48***
853690	−4.94***	−5.05***	−5.75***	−5.40***
854442	−6.39***	−5.91***	−6.72***	−6.17***
870829	−4.66***	−4.40***	−4.89***	−4.67***
870830	−4.30***	−4.41***	−5.28***	−5.05***
870899	−4.51***	−3.72***	−4.68***	−4.07***
903289	−5.68***	−4.97***	−5.52***	−5.07***

Note(s): *** denotes statistical significance at 1% levels
Source(s): Table by authors

Table A1.
Results of the Gengenbach *et al.* (2008) test for cointegration by product

Table A2. Long-run PTM estimates for the bilateral deflated exchange rate ($\widehat{\beta}_i$) and nominal bilateral exchange rate among US dollar and Brazilian real ($\widehat{\beta}_{2,i}$) - by product for the different estimators

	Deflated exchange rate					E_{USD,D_t} ($\beta_{2,i}$)		
	2FE	PMG	MG	CCEP	CCEMG	2FE	CCEP	CCEMG
392690	−0.993	0.018	−1.391	−0.052	−0.334	−0.726	−2.447	−1.500***
401693	−0.077	0.091	0.445	0.055	−1.518***	−2.036***	−1.879*	−1.459***
401699	0.061	−0.035	−0.721	−0.094	−1.185***	−1.472	−1.456	−1.668***
640299	−0.013	0.030	−0.208	0.012	−0.642***	−0.798***	−0.776**	−0.811***
640399	−0.072	0.010	−0.124	0.007	−0.544***	−0.683***	−0.650	−0.657***
731815	−0.643	0.039	−0.083	0.007	−0.909***	−1.097**	−1.016	−1.378***
731816	−1.190**	0.210	−1.793**	0.176	−0.914***	−0.795*	−0.993	−0.947***
731822	−0.236	0.215	−0.937	0.171	−1.454***	−2.167***	−1.551	−1.337***
731829	−0.754	−0.094	−1.368	−0.128	−0.863***	−1.275***	−1.289	−1.164***
732020	0.934	0.048	−0.053	0.046	−1.055***	−2.081***	−1.138**	−1.237***
840991	1.000*	−0.031	0.859	−0.019	−1.014***	−1.217***	−1.037**	−0.985***
840999	0.545	0.030	2.585	0.029	−0.669***	−1.328**	−0.773	−0.809***
841330	0.464**	0.059	0.637	0.040	−0.630***	−1.298***	−0.734	−0.547**
841430	0.231	−0.002	0.688	−0.046	−0.926***	−1.105***	−0.996	−1.077***
842199	−0.555	0.002	−0.644	−0.014	−0.991***	−0.812**	−0.984	−1.004***
848210	−0.228	0.019	0.140	0.012	−0.971***	−0.903***	−1.020**	−1.019***
848310	−1.013	−0.016	−1.379	−0.020	−1.030***	−0.668*	−1.015	−0.859***
848330	−0.504	0.054	−0.458	0.019	−0.899***	−0.582	−1.007***	−1.138***
850152	−0.436	0.003	−0.313	−0.011	−0.756***	−1.241***	−0.912	−0.909***
853690	−0.561	0.016	−1.112	−0.072	−0.662***	−1.219**	−0.728*	−0.627**
854442	0.210	0.129	0.532	0.092	−1.049***	−0.886	−1.152*	−1.061***
870829	0.095	−0.067	2.633	−0.087	−1.033***	−0.973*	−0.966	−1.182***
870830	−1.067*	−0.049	−0.697	−0.095	−0.932***	−1.712**	−1.272	−1.118***
870899	1.147***	−0.076	0.703	−0.112	−0.564***	−0.707	−0.433	−0.517***
903289	−0.635**	0.055	−1.078	0.086	−1.162***	−1.248***	−1.320***	−1.157***

Note(s): *, ** and *** denote statistical significance at 10%, 5% and 1% levels, respectively
Source(s): Table by authors

Variables	Dependent variable $\widehat{\beta}_i$	$\widehat{\beta}_{2,i}$
GDP per capita	0.196**	−0.171*
3926	−0.141	−0.144
4016	−0.280	−0.337
6402	0.411***	0.305
6403	0.521***	0.460**
7318	0.069	−0.040
7320	0.126	−0.003
8409	0.293*	0.255
8413	0.402**	0.524**
8414	0.153	0.046
8421	0.067	0.103
8482	0.128	0.136
8483	0.249	0.291
8501	0.308*	0.204
8536	0.408*	0.431
8544	0.034	0.039
8708	0.302*	0.252
Argentina	0.646***	−0.298
Chile	0.313**	0.071
China	0.184	−0.159
Colombia	0.683***	−0.124
Costa Rica	0.246*	0.065
Dominican Republic	0.516***	−0.099
France	−0.111	0.033
Germany	0.245	0.463**
Guatemala	−0.252	−0.234
Italy	0.044	0.061
Japan	0.122	0.304
Mexico	0.591***	−0.021
Paraguay	0.681***	−0.017
Peru	0.444***	−0.047
South Africa	0.539***	−0.075
Uruguay	0.424***	0.313**
Constant	−3.196***	0.470
Observations	450	450
R-squared	0.214	0.104

Note(s): *, ** and ***, respectively, denote statistical significance at 10%, 5% and 1% levels
Source(s): Table by authors

Table A3. Regression for long run PTM coefficients using estimated dependent variable regression

Supplementary

The supplementary material for this article can be found online.

Corresponding author

Diogo de Prince can be contacted at: dioted@gmail.com