Selection into exporting and international price discrimination: A comparison of Chinese and Indian exporters' pricing behaviour[*]

Sushanta Mallick
Queen Mary University of London, UK
Email: s.k.mallick@qmul.ac.uk

Helena Marques
University of the Balearic Islands, Spain
Email: helena.ferreira-marques@uib.es

ABSTRACT

This paper empirically analyses the export pricing behaviour of Chinese and Indian exporters when there is selection into exporting. Previous pass-through estimates that did not take selection into account are biased if selection into exporting is correlated with the pricing strategy. We use 6-digit product-level data across high- and low-income export destinations over the period 1994-2007 and assess a number of determinants of incomplete pass-through of exchange rates to traded-goods prices, such as the level of external demand, local wage cost, currency volatility and the direction of currency movements. We find systematic differences in the pricing strategies of Chinese and Indian exporters while uncovering a selection bias in exports to high-income markets, although the pricing of exports to low-income markets is independent of the decision to export. Export prices do not increase systematically with the destination market per capita income, and tend to be less sensitive in shipments to advanced nations. Export prices of India are sensitive to the volatility of the trade-weighted effective nominal exchange rate, indicating heterogeneity in prices to maintain competitiveness, but not in China as volatility is insignificant given a fixed currency system. The result is robust to the inclusion of various controls including the development status of the exporter country, a measure of external demand, market share and labour cost. It is also revealed that a country with a relatively flexible currency regime such as India is more likely to exhibit incomplete pass-through due to foreign exchange hedging to minimise the impact of volatility, whereas a country with an inflexible currency system is more likely to have full pass-through as shown in the case of China.

Keywords: *exchange-rate pass-through, pricing-to-market, selection into exporting, product differentiation, India, China*

JEL Classifications: *F4, O1*

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1. Introduction

It has been well-documented that the observed pass-through of exchange rate changes to international prices is incomplete due to sluggish price adjustment originating in mark-up adjustment by the exporters following changes in costs or movements in the exporters' currency (see for example Nakamura and Zerom, 2010, and the references cited there in). The extent to which exchange rate fluctuations affect international prices - exchange rate pass-through (ERPT) - can be affected by the firm's pricing orientation as well as by the degree of exchange rate uncertainty. A substantial literature has documented that exchange rate changes are, at best, weakly associated with changes in traded goods prices at the consumer level (Devereux and Yetman, 2003). Such low degree of pass-through both at the disaggregated level and for aggregate price indexes has been extensively documented (see Mallick and Marques (2012) for the case of India).

However it is important to consider demand in the destination market along with demand in the exporting country, which can be crucial in the price setting behaviour of an exporter. Until now there has been little focus on considering the role of demand-side factors of both exporting and importing countries in explaining price variation across markets. In other words, pricing-to-market (or price discrimination) in response to an exchange rate shock can be conditional on the size of demand in the exporting and importing countries, as measured by their per capita income, which might explain the existence of incomplete exchange rate pass-through (Ferrantino, Feinberg and Deason, 2012). Besides, Hoffmann (2007) shows that there are significant differences in the variability of macroeconomic aggregates under fixed and flexible exchange rate regimes.

Finally, price discrimination causes ERPT to be incomplete in both the short and the long run with high exchange-rate volatility (Corsetti *et al*., 2008). It is therefore of interest to compare the pricing behaviour in countries with fixed exchange rate regimes, such as China, to countries with flexible exchange rate regimes, such as India. In this context, considering two key emerging market exporters (China and India), where exchange rate fluctuations are respectively fully and partially managed by the authorities of these two countries, can reveal whether exchange rate volatility tends to increase price discrimination and thereby reduce the degree of pass-through. The interest of the comparison is augmented by the fact that the two countries considered in this study are both important emerging markets that under the current economic downturn have taken up the role of growth engines in the world economy.

The explanations given in the literature for the low pass-through that is commonly found are primarily microeconomic, such as pricing to market (PTM) by imperfectly competitive firms (Corsetti and Dedola, 2005). Imperfect competition in international markets is one of the key explanations for why prices of 'similar' goods might differ among destination markets (Aw, 1993). Aw *et al*. (2001) isolate the market-specific price differences by simultaneously accounting for firm-level price heterogeneity in the same product market. If firms face capacity constraints in distribution networks or quantitative trade restrictions, then pricing-to-market may be greater during depreciations of the exporter's currency, or if firms attempt to build market share subject to the threat of trade restrictions, then pricing-to-market may be greater during appreciations of the exporter's currency (Knetter, 1994). This asymmetry in pass-through is also studied at both the firm and the aggregate levels by Rodriguez-Lopez (2011). Assuming a home currency depreciation, he finds that, when firms have heterogeneous productivity, aggregate exchange rate pass-through into home import prices can be negative even if at the firm level the pass-through is positive (although incomplete). This result is due to the adjustment of the extensive margin whereby only the most productive foreign exporters survive a depreciation of the home currency and each exporter adjusts the mark up differently depending on productivity.

In this paper we compare the pricing-to-market decisions of two major emerging exporters (China and India) in response to changes in their NEER (Nominal Effective Exchange Rate). The results suggest that exporters absorb changes in the Indian rupee or the Chinese yuan by changing export prices in their own currency in the opposite direction to that of the exchange rate change. We further establish whether the decision to stay in or out of an export market for a particular product is correlated to pricing-to-market decisions. If this correlation exists but it is not taken into account, the estimates of exchange rate pass-through will be biased. The results indicate that a selection bias exists in exports to high-income markets, but the pricing of exports to low-income markets is independent of the decision to export. However, we find no evidence that firms that enter and exit are more likely to be sensitive to exchange rate changes when the exchange rate appreciates than when it depreciates.

We further find that Indian exporters have been able to vary mark-ups over different markets to stay competitive in response to the rupee's volatility, while Chinese exporters do not react to currency volatility given a fixed exchange rate system with a narrow band. This suggests that Indian exporters exercise market power to obtain price premia when their currency experiences volatility with respect to the importer's currency. There is a strong negative relationship between volatility and pass-through in the case of a (relatively) flexible currency system (India) implying low pass-through, while high pass-through occurs in the case of a fixed exchange rate regime (China). We also find that PTM is largely symmetric: exporters absorb appreciations of their currency through downward price adjustment while they tend to increase prices following depreciation of their currency. The resulting pass-through is incomplete in India's exports to high-income markets. There is no pass through in the remaining cases. The results are very robust across specifications.

Section 2 of the paper develops a theoretical model from which the empirical specification is derived. Section 3 deals with estimation and data issues, section 4 presents the estimation results with several robustness checks, and section 5 concludes.

2. The model

We outline a simple model of exchange rate pass-through in a similar spirit as Devereux and Yetman (2003), Ghironi and Melitz (2005), Melitz and Ottavano (2008), Chaney (2008) and Rodriguez-Lopez (2011). In this class of models based on the work of Melitz (2003), it is assumed that only a subset of domestic firms are exporters due to the interplay between heterogeneous productivity across firms and, in some models, the existence of fixed costs of exporting.

A firm located in country *i* and exporting to country *j* faces marginal and fixed costs in terms of domestic currency and sets prices in terms of domestic currency. The demand faced by the exporter in the overseas market is given by

$$C_{ij} = \left(\frac{P_j^*}{p_{ij}^*} \right)^{\lambda} C_j \qquad [1]$$

where p_{ij}^* is the firm's price of its exports to the destination market given in foreign currency, P_j^* is the composite price index for all foreign goods sold on the destination market, also given in foreign currency, C_j is the expenditure level, or absorption, of the destination market; and λ is the price elasticity of external market demand, which is country-specific and a function of the exchange rate (see Corsetti and Dedola, 2005). This type of demand function is usually derived from the domestic country's utility maximisation (see Betts and Devereux, 2000 or Helpman et al, 2008). The exporting firm gets a share of the destination market that depends on its price relative to the composite price index that includes the prices of all sellers.

Furthermore, the firm's price in foreign currency is obtained from its price in domestic currency by means of $p_{ij}^* = e_{ij} p_{ij}$, where e_{ij} is the bilateral exchange rate defined as the units of foreign currency per unit of domestic currency, such that an increase in the exchange rate means an appreciation. The composite price index in the foreign market can also be converted to domestic currency by the same means.

Following the formulation in Chaney (2008), the exporting firm's profit in terms of domestic currency is given by:

$$\pi_{ij} = \left(p_{ij}^{*} - \frac{w_{it}\tau_{ij}}{\varphi_i} \right) C_{ij} - F_{ij} \qquad [2]$$

where $\frac{w_i}{\varphi_i}$ is the productivity-adjusted wage cost at the producer's location, τ_{ij} is the iceberg transport cost which depends on distance, and F_{ij} is the fixed cost of exporting, which is country-specific but not firm-specific. Thus the profit-maximization problem faced by a firm in an imperfectly competitive industry can be derived by maximizing profit with respect to the choice variable, p_{ij}. The first order condition can be written as:

$$f'\left(C_{ij}\right)\left(p_{ij}^{*} - \frac{w_i\tau_{ij}}{\varphi_i} \right) = -C_{ij} \qquad [3]$$

Substituting the demand function [1] in this first order condition and assuming that the exporting firm could adjust its price at any time, the equilibrium export price can be derived as:

$$p_{ij} = \frac{\lambda}{\lambda - 1} \frac{e_{ij} w_i \tau_{ij}}{\varphi_i} \qquad [4]$$

This pricing equation is a mark-up equation modified to reflect the existence of transport costs and heterogeneous firm productivities. Whilst wages and transport costs are defined at the country-level, productivity is defined at the firm level. This presents us with the problem of obtaining firm-level data and carry out the empirical work at the firm-level, as has been done by Chaney (2008) for the US, Berman et al (2012) for France, Manova and Zhang (2012) for China, Chatterjee et al (2010) for Brazil, among others. In the absence of consistent cross-country firm-level datasets, an alternative approach – which allows the use of product-level data whilst proxying for the unobservable φ_i – has been proposed by Helpman et al (2008).

Firms self-select into export markets according to their productivity level, such that there is a cut-off value $\overline{\varphi}_i$ defined by the firm's zero-profit condition which determines the fraction of firms exporting to the destination market. That cut-off value can be shown to be equal to:

$$\overline{\varphi}_i = \left(\frac{F_{ij}}{C_j}\right)^{1-\lambda} \left(P_j^*\right)^{\frac{\lambda}{1-\lambda}} w_i \tau_{ij} \varepsilon_{ij} \quad [5]$$

where $\varepsilon_{ij} = \left(\frac{\lambda}{\lambda-1} e_{ij} - 1\right)^{\frac{1}{1-\lambda}} \left(\frac{\lambda}{\lambda-1} e_{ij}^2\right)^{\frac{-\lambda}{1-\lambda}}$ collects the bilateral exchange rate terms, with its derivative with respect to the bilateral exchange rate depending on the price elasticity of demand.

Thus, as in Helpman et al (2008), the fraction of firms exporting to the destination market is a function of exporting country variables (w_i), importing country variables (P_j^*, C_j) and country-pair variables $(\tau_{ij}, F_{ij}, e_{ij})$. Accordingly, they suggest that the productivity cut-off, and consequently the associated probability of exporting, can be estimated in a Probit model as a function of exporting country variables, importing country variables and country-pair variables. In this way we can take into account the effect of firm heterogeneity on pricing using only product-level data. The log-linear approximation of equation [5] can be written as:

$$\ln \bar{\varphi}_i = \gamma_0 + \gamma_i + \gamma_j + \gamma_{ij} \quad [6]$$

where $\gamma_i = \ln(w_i)$, $\gamma_j = \frac{\lambda}{1-\lambda}\ln(P_j^*) - (1-\lambda)\ln(C_j)$ and

$\gamma_{ij} = (1-\lambda)\ln(F_{ij}) + \ln(\tau_{ij}) + \ln(\varepsilon_{ij})$.

In this setup, the conditional probability of exporting (ρ_{ij}) is defined as:

$$\rho_{ij} = \Pr\left(T_{ij} = 1 \mid \text{observed variables}\right) = \Phi\left(\ln \bar{\varphi}_i\right)$$

where T_{ij} is an indicator variable that takes value 1 when there is an export flow from country *i* to country *j* (prices are observed) and takes value 0 otherwise (prices are not observed).

In a second stage, the predicted probabilities are included in the estimation of the pricing equation in order to obtain consistent estimates of exchange rate pass-through given that firms are heterogeneous and not all select into exporting, this is, the share of firms that export and the selection of product-market pairs into trade both vary. The log-linear approximation of the pricing equation [4] is:

$$\ln p_{ij} = \ln\left(\frac{\lambda}{\lambda - 1}\right) + \ln\left(e_{ij}\right) + \ln w_i + \ln \tau_{ij} - \ln \hat{\varphi}_i \quad [7]$$

Facing different demand levels in each market, the exporting firm will establish a market-varying mark-up over marginal costs. The mark-up established over destination country *j* partly depends on the wage level of that country (Alessandria and Kaboski, 2011). If we assumed, as in Rodriguez-Lopez (2011), that wages are sticky, we could proxy wages by an exporting country specific fixed effect. However, his model is a developed country model, whereas here we are working with emerging markets showing very fast growth, including wage growth. Indeed, Chamarbagwala and Sharma (2011) argue that industrial wages have risen in India due to the labour force's skill upgrading, among other factors. Therefore, it would not be appropriate to assume sticky wages. In a context of very fast growth, it is preferable to assume almost perfect sectoral mobility within each country and to use the average manufacturing wage in each country as a measure of production costs. In this way, we are still able to capture time variation in those costs. In work on other emerging markets, Alvarez and Fuentes (2011) use the income per capita of Chile's export markets, whilst Marmolejo (2011) includes both Mexican and US wages in a model of exchange rate pass-through into Mexican import prices after the constitution of NAFTA. In the absence of wage data, using income per capita would be a good proxy to control for increasing globalisation of production activity, when a large share of international trade occurs through intra-firm transactions, leading to incomplete pass-through (see Hellerstein and Villas-Boas, 2010). Ferrantino, Feinberg and Deason (2012) have also used the per capita income of exporters to introduce vertical differentiation and the per capita income of importers to introduce pricing-to-market in a cross-section of 6-digit unit values for 2005. We will start by using income per capita as data is available for the whole sample and later check for the robustness of the results using wage data.

The lack of responsiveness of export pricing to exchange rate fluctuations may be partially on the back of hedging activities by trading agents due to foreign exchange volatility to eliminate exchange rate risk, as hedging against foreign exchange uncertainty can affect the structure of pricing behaviour and pass-through. In addition to the variables reflecting mark-up adjustment, a reduction in currency risk exposure due to hedging activities could lead to a decline in the transmission of changes in the exchange rate. Thus pricing-to-market estimates must be obtained by controlling for bilateral foreign exchange volatility in order to observe whether there is a differential impact of high and low volatile destination markets on international pricing.

3. Empirical specification and data

Using a log-linear version of equation [7] with discrete change, the following estimable equation can be written:

$$\begin{aligned}\Delta \ln p_{ij,t}^{k} = {} & \beta_0 + \beta_1 \Delta \ln\left(e_{ij,t-1}\right) + \beta_2 \ln GDPpc_{i,t-1} + \beta_3 \ln GDPpc_{j,t-1} + \\ & + \beta_4 \operatorname{var}\left[\Delta \ln\left(e_{ij,t-1}\right)\right] + \beta_5 pshare_{ij,t-1} + \beta_6 HSshare_{i,t-1}^{k} + u_{ij,t}^{k}\end{aligned} \qquad [8]$$

where $GDPpc_i$ and $GDPpc_j$ are the exporter and the importer GDP per capita (later replaced by sectoral wages to check robustness) and *e* is the exporting country's NEER with a rise indicating an appreciation of the exporter's currency. Beladi *et al*. (2010) develop a model of exchange rate pass-through allowing for a stochastic process of the exchange rate. Here we capture that stochastic process by including a lagged exchange rate variable .

Also note that $\beta_0 = \ln\left(\frac{\lambda}{\lambda-1}\right)$. Therefore the constant term gives us information about the price elasticity of external market demand for each market in each moment in time. This elasticity determines the base price level. On the other hand, the pricing equation's error term *u* is correlated with unobservable firm-level productivity. If this variable non-randomly determines self-selection into export markets, the estimation of a selection model will correct any selection bias in the estimates of the pricing equation.

We take three variables as measures of trade costs τ_{ij}. The first measure is exchange rate volatility, specifically currency risk expressed as $\mathrm{var}\left[\Delta \ln\left(e_{ij,t-1}\right)\right]$, which may explain why markets have not become fully integrated, as in the case of deviation from absolute PPP as evidenced in Alessandria and Kaboski (2011). Hedging is one such activity that aims to reduce trade costs and hence we need to control for this factor before deriving the PTM or ERPT estimates. If these hedging activities are not taken into account, the average pass-through coefficient could be underestimated. If the estimated degree of pass-through is used to measure the market or pricing power, such power in the industry may also have been underestimated without considering the impact of exchange rate volatility. Our measure of exchange rate volatility is obtained according to the procedure explained in Mallick and Marques (2010). Briefly, we use a GARCH (1, 1) model for variance as the simplest and most robust of the family of volatility models which looks like this:

$$h_t = \omega + \alpha h_{t-} \varepsilon_{t-} + \beta h_{t-} \; .$$

This model computes the variance (h) of the current exchange rate as a weighted average of a constant and previous period's variance forecast and squared error.

The two other measures of trade costs, or the costs of exporting, are the share of exporter i in market j $\left(pshare_{ij,t-1}\right)$ and the share of product k in exporter i's export basket $\left(HSshare_{i,t-1}^{k}\right)$. As in Helpman et al (2008), we consider that a high presence in a destination market or in a product market lowers the costs of exporting to that country or of exporting that product. Furthermore, the productivity effect is revealed indirectly through a two-stage estimation procedure. As in Helpman et al (2008), the productivity cut-off, and consequently the associated probability of exporting, will be estimated in a Probit model as a function of exporting country variables, importing country variables and country-pair variables. In this way we can take into account the effect of firm heterogeneity on pricing using only product-level data.

Later we also consider asymmetry as a robustness check. The existence of asymmetry means that exchange rate risk (volatility) affects exports differently during appreciations and depreciations, which may reflect the exporter's asymmetric risk perception and hedging behaviour (Fang *et al*., 2009). We can identify how the β_1 coefficient in equation (8) differs when the domestic currency value declines (more entry of exporting firms), and when the currency appreciates (more exit of exporting firms). To do this, we consider the exchange rate separately when it appreciates and when it depreciates. Figure 1 shows that the NEER of China has appreciated until 2001, depreciated in 2002-05 and appreciated again from then on. The NEER of India has fluctuated around a depreciating trend until 2006 and appreciated in the last sample year.

Figure 1 here

We employ a panel data set from UN Comtrade consisting of location- and product-specific export price data from China and India in 1994-2007 to show the relative market power of Chinese and Indian exporters in different product categories during our sample period, allowing us to identify price discrimination or PTM in traded goods at the 6-digit level. Given the global crisis that has been unfolding since 2008 which has interfered with the normal trade flows due to lack of credit to firms, we use data up to 2007, distinguishing Chinese and Indian exports into high and low-income countries. Table A1 in the Appendix shows export price data availability for our sample of high and low-income markets, which are the main markets of China and India. As this ranking does not fully coincide for the two countries, later we check the robustness of our results using only the markets that are common to the ranking of both exporters. NEER data is taken from IMF IFS (2005=100). GDP per capita is taken from the World Bank Development Indicators.

Our main objective is to compare the PTM of China and India exports to high and low-income markets. The PTM coefficient would be higher in high-income markets due to consumer search behaviour determined by higher wages. Hence, although the share of emerging countries in high-income markets has declined (see Figure 2), we would still expect their PTM to increase in advanced markets whereas relative PPP could hold in the case of low-income markets (see Figures 3-3a). Table A2 in the Appendix shows that the average export price of China to high-income markets is about twice as high as to low-income markets, however in the case of India the reverse occurs. On the other hand, Table A3 in the Appendix reveals that the average export price distribution at the product level is similar for China and India, although which country shows higher average price within each product group varies depending on particular categories. Comparing China and India from the point of view of product composition will reveal the product categories where they face tougher competition. China is very strong in labour-intensive sectors but India fares well in some high-tech sectors. We examine how this gets reflected in the PTM coefficient into their major markets.

Figures 2, 3, 3a here

4. Estimation results

4.1. Benchmark results

Due to the very large number of observations (over 1 million), we start with a simple panel regression that does not require great computational capacity. The Hausman test carried out for fixed and random effects reveals that the random effects estimator is inconsistent. Hence we prefer the consistent, although less efficient, fixed effects estimator. Table 1 presents a version of this estimation using variance-covariance estimates clustered by exporter-importer-product groups to account for correlation of observations within each group. This essentially recognizes that exports of the same product to the same market are correlated over time and accounts for some persistence in export patterns.

Table 1 here

Although there is foreign exchange intervention in the two countries, the yuan has been managed more strictly than the rupee, so exchange rate volatility is not significant for China in high-income markets. It does increase export prices in the case of India and also for China in low-income markets, as these markets' currencies are themselves more volatile. After accounting for exchange rate volatility, the results reveal a very different pricing behaviour by Chinese and Indian exporters. Chinese export prices change on average 15% more than the exchange rate (13% in high-income markets and 20.5% in low-income markets), in effect amplifying exchange rate changes in the opposite direction: local currency prices decrease with appreciations and increase with depreciations. From a visual inspection of Figure 4, one can conclude that there is a similar stabilizing price effect in response to currency appreciation in the case of China and India. However, in the case of India's export prices, we cannot discard the possibility of a one-to-one price change with respect to the exchange rate, so that when there is an appreciation (depreciation) producer currency prices decrease (increase) to the same extent, leaving local currency prices unchanged (zero pass-through). Thus Indian exporters seem to be practicing classical local currency pricing, whereas Chinese exporters overreact to exchange rate changes to an extent that the exchange rate fluctuations become destabilizing.

Figure 4 here

We also observe that producer currency prices increase with the exporter's income in all cases, but react positively to the importer's income in the case of China's exports to low-income markets and negatively in the case of India's exports to high-income markets. Hence, following the interpretation of Ferrantino, Feinberg and Deason (2012), who found significant exporter-income effects and importer-income effects for 56% of the products in their sample, in our case China exports higher-quality goods to its low-income markets, and India exports lower-quality goods to its high-income markets. Moreover, product and market shares, by lowering the costs of exporting, reduce producer currency prices. Finally, the constant term (price elasticity of external market demand) impacts negatively on export prices, as would be expected, but has no effect on exports to high-income countries, which could be due to the nature of the products being exported to these markets (quality, differentiation).

4.2. Selection issues

The sole estimation of the pricing equation does not, however, take into account that zero or missing unit values do not happen by chance. Zeros or missing values can be found because exporters do not self-select into exporting certain goods to certain markets. To account for this, we introduce a Heckman (1976) procedure in the estimation of ERPT. We consider that self-selection into exporting a certain product to a certain market is a function of market characteristics such as the income level of the market, of exporter characteristics such as its share in the market and of product characteristics such as its share in the exporter's export basket. Given that an exporter decides to export a certain product to a certain market in the first stage, these variables also determine the product's price in that market in the second stage. If the unobservable factors determining both self-selection into an export market and the price placed to that market are correlated (we will test for this), the benchmark results would be biased. The Heckman procedure returns consistent and asymptotically efficient estimates in such cases.

The empirical strategy consists on estimating regression equation [8] assuming that the dependent variable $\Delta \ln p_{ij,t}^{k}$ is observed when

$$\gamma_0 + \gamma_1 \ln GDPpc_{j,t-1} + \gamma_2 pshare_{ij,t-1} + \gamma_3 HSshare_{i,t-1}^{k} + v_{ij,t}^{k} > 0$$

where the pricing equation error $u_{ij,t}^{k}$ and the selection equation error $v_{ij,t}^{k}$ have correlation equal to ρ. Here we consider that an exporter would decide whether to export to an importer-product pair based on the importer's income, as well as the exporter's share in that market and the share of the product in the exporter's export basket as measures of the cost of exporting.

The two-step model is generally stable and tolerates estimates of ρ outside the -1 and 1 correlation bounds. For these reasons, the two-step model may be preferred when exploring a large dataset. Still, if the maximum likelihood estimates cannot converge, or converge to a value of ρ that is at the boundary of acceptable values, there is usually no support for fitting a Heckman selection model to the data.

The results for the consistent and efficient two-stage Heckman estimator are shown in Table 2. In our sample, most zeros and missing values are found in exports to high-income countries, implying a narrower product range (smaller extensive margin) in exports to those markets. This effect is more pronounced for China than for India, meaning that India's exports are more diversified, and possibly of higher quality or higher technological content, than those of China. In addition, the two-stage selection model is justified only in the case of China's exports to high-income countries, as in the remaining cases there is no evidence of significant correlation between the residuals of the selection equation and those of the regression equation. In the absence of a significant correlation between those residuals, we can consider the benchmark results as unbiased. However, for the case of China's exports to high-income markets, there is a significant (positive) correlation between the unobservable factors determining both the export and the pricing decisions. In this case we find that products with a higher share in China's export basket have a higher probability of being exported to high-income markets, but the importer's income, China's share in the importer's market and the price elasticity of demand in the importer's market decrease both the probability of exports from China and the price of the products being exported. These results imply that China's exports to high-income markets are essentially mature products with little differentiation.

Table 2 here

In Table 3 we present an alternative set of results using a pseudo-likelihood estimator that allows the clustered estimation of the standard errors of the Heckman selection model by importer-product pairs. This estimator converges slowly in large datasets and may not be efficient (the robust standard errors may be a bit larger than those obtained in the two-stage procedure). However, the clustering of standard errors renders the results closer to the benchmark, which is unbiased in most cases. Any systematic differences between the two-stage estimator and the pseudo-likelihood estimator with clustered standard errors are due to importer-product-specific correlations that the two-stage model does not take into account. Moreover, the hypothesis of independent equations is accepted only in the case of exports to low-income countries, so the selection model is justified in the case of exports to high-income countries.

Table 3 here

4.3. Common markets

We re-estimate equation [8] for a subsample of markets common to both exporters (China and India). In our sample, there are 8 common high-income markets and 16 common low-income markets (see Table A1 for detailed information). In this way, we make sure that export market selection is not driving differences in results. These results are shown in Table 4 and broadly preserve the benchmark features, except that China's export prices to high-income countries now respond negatively to China's income (not responding to exchange rate volatility) and India's export prices to high-income countries also respond negatively to the importers' income. Moreover, for this subsample of markets Indian exporters practise incomplete pass-through, absorbing an average of 59% of the exchange rate change (45% for high-income markets and 80% for low-income markets). This difference in pass-through may be caused by the export basket composition, with more differentiated goods with less elastic demand being exported to high-income countries and more homogeneous goods with more elastic demand being exported to low-income countries.

Table 4 here

Since the Heckman selection model has been shown to be adequate in some cases, we have checked those results using the pseudo-likelihood with robust standard errors version (not shown, but available). The selection model confirms (at the 10% significance level) that the prices of China's exports to high-income countries respond negatively to China's income (anyhow the hypothesis of independent equations is accepted). However, in the case of India, the income of importing countries is no longer significant (with the hypothesis of independent equations being rejected at the 5% significance level). Indeed, the importer's income does influence (negatively) the probability of exporting by India's exporters, so we can consider the importer's income coefficient biased in the fixed effects model.

Another important bias in the fixed effects model is an upward bias (in absolute value) of the product share coefficient. This is because the fixed effects model does not distinguish between the pricing and exporting decisions and the selection model shows that the product's share in the exporter's export basket is significant for the export decision in most cases. This variable has been taken as a (reverse) proxy for the costs of exporting: when the product's share in the exporter's export basket is high, the export channels for that product are already established and the export know-how already exists, so entering an additional market in that situation may have lower costs than when the product is seldom exported. Therefore export prices could have a downward trend in response to higher market shares as shown in Figure 5.

Figure 5 here

4.4. Testing for asymmetry

We extend equation [8] to distinguish between changes in the lagged log of the NEER under depreciation (D) and under appreciation (A):

$$\Delta \ln p_{ij,t}^{k} = \beta_0 + \beta_1^{D}\Delta^{+} \ln\left(e_{ij,t-1}\right) + \beta_1^{A}\Delta^{-} \ln\left(e_{ij,t-1}\right) + \beta_2 \ln GDPpc_{i,t-1} + \beta_3 \ln GDPpc_{j,t-1} + \\ + \beta_4 \operatorname{var}\left[\Delta \ln\left(e_{ij,t-1}\right)\right] + \beta_5 pshare_{ij,t-1} + \beta_6 HSshare_{i,t-1}^{k} + u_{ij,t}^{k} \quad [9]$$

In our sample, appreciation happens in 2/3 of the observations (around 1 million). This still leaves half a million of observations for which depreciation takes place. We can identify how the β_1 coefficient in equation (9) differs when the value of domestic currency depreciates (net entry of exporting firms), and when it appreciates (net exit of exporting firms). The existence of asymmetry means that exchange rate risk (volatility) affects exports differently during appreciations and depreciations, which may reflect the exporter's asymmetric risk perception and hedging behaviour (Fang *et al*., 2009).

The fixed effect results are shown in Table 5. Overall, the results follow the benchmark and it is not possible to detect any asymmetry effects (the 95% confidence intervals for the exchange rate coefficients under appreciations and depreciations always overlap). However, the consideration of asymmetry reveals that for China we cannot discard the possibility of a one-to-one price change with respect to the exchange rate, so that when there is an appreciation (depreciation) producer currency prices decrease (increase) to the same extent, leaving local currency prices unchanged (zero pass-through). We also find zero pass-through in the prices of India's exports to low-income markets, with exporters fully absorbing exchange rate changes. On the contrary, for the case of India's exports to high-income markets, the producer currency price change absorbs around 44% of the exchange rate change, so that there is incomplete pass-through.

Table 5 here

4.5. Using industrial wages

We have used GDP per capita data as consistent and comparable series were available for all sample countries. However, it would be preferable to use wages of industrial sectors as these reflect more directly the industry's production costs. These are not available for all sample countries at the sectorial level and are given in different units (month, week or day) which have to be converted to hourly rates. Therefore the wage data contains substantial measurement error. Nevertheless we present those results in Table 6 as a robustness check. They do not generally differ from previous tables and do not alter the conclusions of the paper.

Table 6 here

5. Conclusions

The conventional thinking has long been that ERPT is always complete and rapid in developing economies, as they are price takers and hence cannot exercise PTM. In this paper, we find diverse pricing strategies at a 6-digit product level for Chinese and Indian exporters. The paper investigated the degree of PTM or the pricing behaviour of Chinese and Indian exporters across destination markets, controlling for the destination market per capita income, the currency volatility of the exporter and uncovering any asymmetric pattern in price variation. Considering export prices from two countries with different exchange rate regimes, we are able to explore the effect of both fixed and flexible exchange rate regimes on price variations. We show that pass-through has been higher and faster for a country with a fixed regime (China) relative to a country with a managed floating currency regime (India).

This paper presented a new analysis of the sources of incomplete pass-through, uncovering the existence of a selection bias in exports to high-income markets due to a narrow extensive margin of exports by emerging market exporters and showing that Indian exporters are more sensitive to exchange rate changes in high-income markets, while in low-income markets they tend to balance their market shares with increasing their mark-ups. Thus we conclude that external demand conditions and the degree of currency volatility play an important role in relating exchange rate changes to price variations in the buyers' currency, in this way establishing the evidence of differences in PTM behaviour of China and India across their most important high- and low-income export destinations.

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Figures and Tables

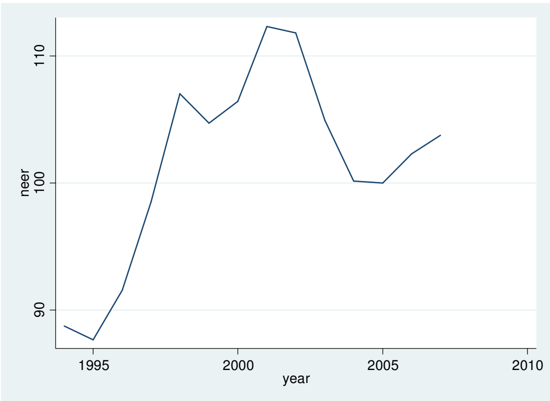
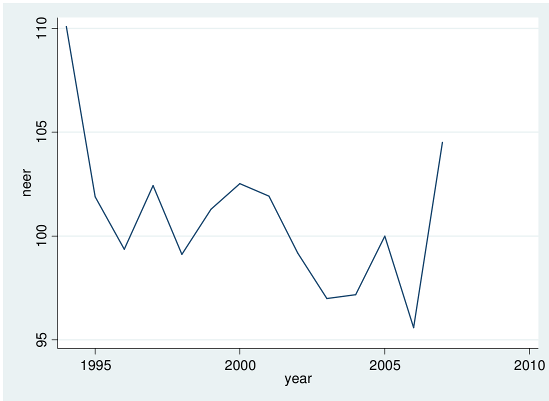
Figure 1: NEER in China and India (1994-2007)	
China	India
	

Figure 2: Shares in export market by China and India (1994-2007)

Figure 3: Unit value of exports for China and India's exports

Figure 3a: Log of unit value of total exports by China and India

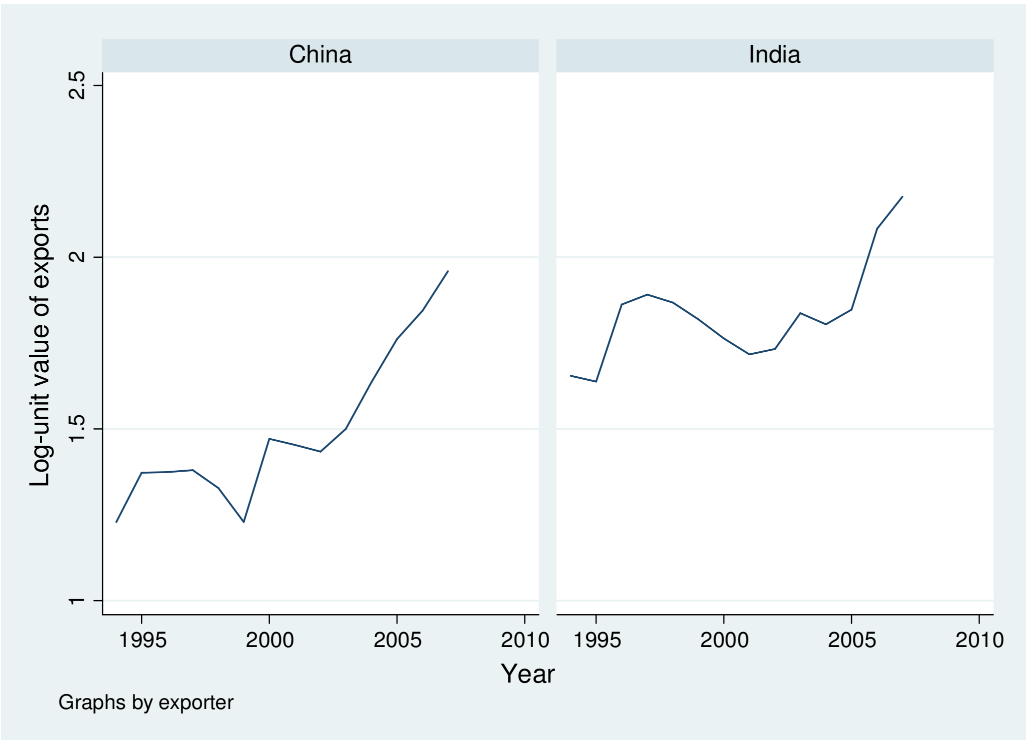


Figure 4: Log of Unit value of exports for China and India against NEER

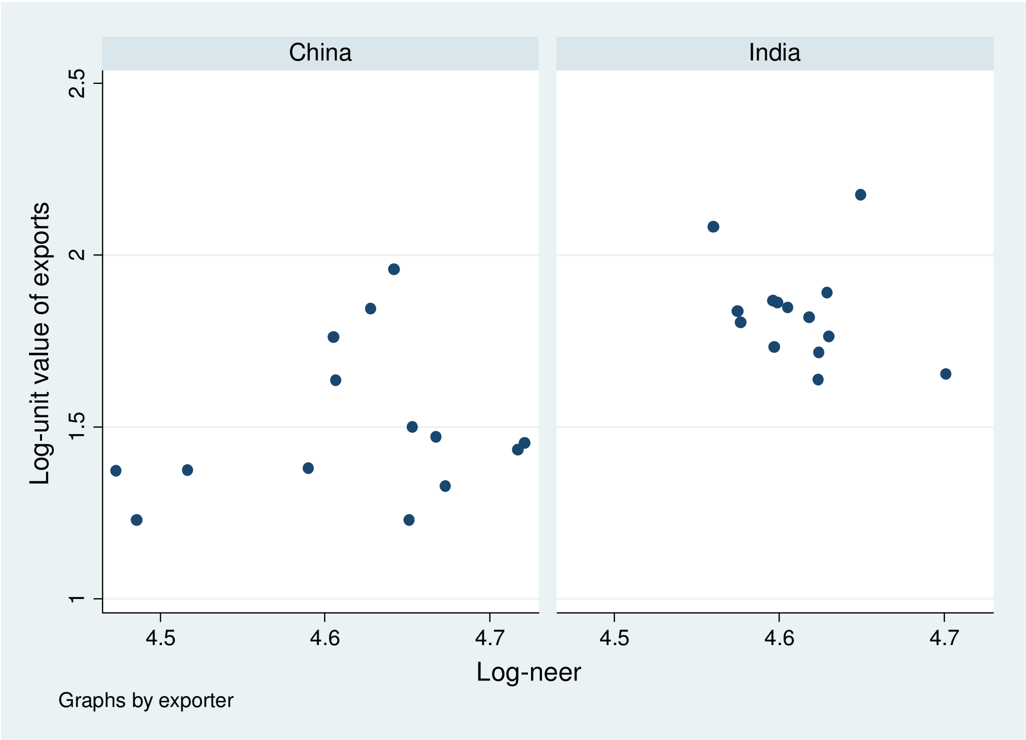


Figure 5: Log of Unit value of exports for China and India against their market share

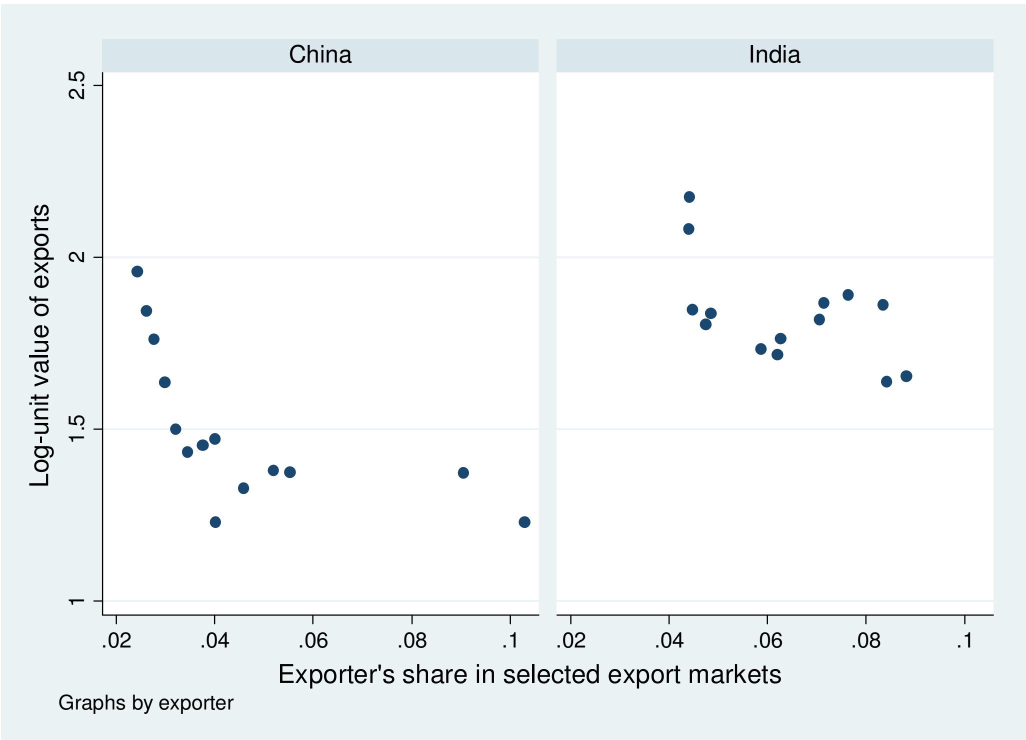


Table 1: panel fixed effects							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	all	China	India	China to high_inc	China to low_inc	India to high_inc	India to low_inc
$\Delta \ln\left(e_{ij,t-1}\right)$	-1.197**† (0.024)	-1.151**† (0.027)	-0.740**† (0.102)	-1.133**† (0.033)	-1.205**† (0.045)	-0.758** (0.130)	-0.760** (0.163)
$\ln GDPpc_{i,t-1}$	0.047** (0.004)	0.017** (0.005)	0.117** (0.009)	0.016** (0.006)	0.031** (0.008)	0.137** (0.014)	0.122** (0.014)
$\ln GDPpc_{j,t-1}$	0.048** (0.006)	0.063** (0.007)	-0.012 (0.012)	0.009 (0.010)	0.087** (0.008)	-0.060** (0.021)	0.004 (0.014)
$\text{var}\left[\Delta \ln\left(e_{ij,t-1}\right)\right]$	17.229** (1.194)	2.685 (1.828)	30.406** (2.359)	-1.436 (2.313)	6.682* (2.986)	35.264** (2.969)	19.831** (3.890)
$pshare_{ij,t-1}$	-0.313** (0.020)	-0.425** (0.026)	-0.146** (0.032)	-0.337** (0.026)	-1.069** (0.100)	-0.152** (0.035)	-0.094 (0.079)
$HSshare_{i,t-1}^{k}$	-2.813** (0.365)	-3.736** (0.672)	-1.925** (0.415)	-4.106** (0.927)	-3.474** (0.876)	-3.772** (1.189)	-1.711** (0.459)
Constant	-0.733** (0.039)	-0.642** (0.049)	-0.632** (0.074)	-0.130 (0.085)	-0.861** (0.056)	-0.250 (0.147)	-0.797** (0.079)
Observations	948891	691488	257403	385500	305988	149859	107544
Importer-product groups	154253	107195	47058	55293	54327	23556	23502
F-test	988.40**	799.83**	228.07**	461.06**	331.99**	165.22**	75.37**
Robust standard errors in parentheses. * significant at 5%; ** significant at 1%. † different from 1 at 5%.							

Table 2: Heckman selection model -- two-step estimates							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	all	China	India	China to high_inc	China to low_inc	India to high_inc	India to low_inc
Regression equation (pricing decision)							
$\Delta \ln\left(e_{ij,t-1}\right)$	-1.219**† (0.023)	-1.096**† (0.028)	-0.754**† (0.085)	-1.122**† (0.035)	-1.092*† (0.519)	-0.794**† (0.110)	-0.737 (0.703)
$\ln GDPpc_{i,t-1}$	0.037** (0.002)	0.033** (0.004)	0.103** (0.007)	0.013** (0.005)	0.059 (0.070)	0.098** (0.010)	0.119* (0.059)
$\ln GDPpc_{j,t-1}$	-0.002 (0.003)	-0.002 (0.002)	0.001 (0.003)	-0.082** (0.007)	0.053 (0.123)	-0.009 (0.015)	-0.018 (0.074)
$\text{var}\left[\Delta \ln\left(e_{ij,t-1}\right)\right]$	15.836** (1.004)	0.770 (1.866)	29.049** (1.782)	-2.503 (2.365)	3.593 (35.324)	34.172** (2.226)	18.839 (15.811)
$pshare_{ij,t-1}$	-0.175** (0.032)	-0.184** (0.028)	-0.066** (0.022)	-0.151** (0.014)	0.278 (1.828)	-0.075** (0.023)	0.440 (2.009)
$HSshare_{i,t-1}^{k}$	-0.907** (0.173)	-1.064** (0.263)	-1.279 (0.738)	-0.889** (0.324)	-3.763 (8.609)	-2.037 (1.236)	-0.919 (2.563)
Constant	-0.239** (0.024)	-0.191** (0.033)	-0.742** (0.068)	0.763** (0.079)	-1.257 (2.210)	-0.576** (0.119)	-1.192 (1.948)
Selection equation (exporting decision)							
$\ln GDPpc_{j,t-1}$	-0.215** (0.002)	-0.341** (0.003)	-0.015** (0.003)	-1.547** (0.011)	0.047** (0.007)	-0.149** (0.019)	-0.016* (0.009)
$pshare_{ij,t-1}$	-1.147** (0.014)	-1.378** (0.015)	-0.105** (0.035)	-1.353** (0.016)	0.890** (0.171)	-0.177** (0.038)	0.548** (0.104)
$HSshare_{i,t-1}^{k}$	-1.874** (0.239)	12.958** (1.226)	-3.305** (0.272)	24.420** (1.889)	-2.292* (1.106)	-6.538** (0.569)	-0.492 (0.524)
Constant	3.754** (0.020)	4.989** (0.029)	1.818** (0.030)	17.448** (0.112)	1.778** (0.052)	3.198** (0.199)	1.794** (0.069)
Uncensored observations	948891	691488	257403	385500	305988	149859	107544
Censored observations	52205	39211	12994	34347	4864	7755	5239
Wald Chi2-test	3637.72**	2697.47**	1197.01**	1633.60**	9.37	905.57**	12.06*
Mills λ	0.289**	0.226**	0.733	0.255**	11.221	0.411	5.563

* significant at 5%; ** significant at 1%. † different from 1 at 5%.
The selection effect is summarized by $\lambda = \rho * \sigma$ (with ρ the correlation between the residuals of the regression equation and the selection equation and σ the standard error of the residual in the regression equation). A significant λ justifies the use of a selection model.

Table 3: Heckman selection model – pseudolikelihood estimates							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	all	China	India	China to high_inc	China to low_inc	India to high_inc	India to low_inc
Regression equation (pricing decision)							
$\Delta \ln\left(e_{ij,t-1}\right)$	-1.218**† (0.021)	-1.095**† (0.026)	-0.755**† (0.099)	-1.117**† (0.032)	-1.090** (0.042)	-0.795** (0.086)	-0.737** (0.156)
$\ln GDPpc_{i,t-1}$	0.038** (0.001)	0.033** (0.004)	0.103** (0.005)	0.016** (0.005)	0.060** (0.006)	0.098** (0.007)	0.120** (0.008)
$\ln GDPpc_{j,t-1}$	0.001** (0.001)	-0.001 (0.001)	0.003** (0.001)	-0.069** (0.005)	0.007** (0.001)	0.0001 (0.007)	-0.002 (0.003)
$\text{var}\left[\Delta \ln\left(e_{ij,t-1}\right)\right]$	15.805** (1.118)	0.818 (1.789)	29.033** (2.275)	-2.383 (2.274)	3.484 (2.890)	34.155** (2.899)	18.821** (3.684)
$pshare_{ij,t-1}$	-0.133** (0.007)	-0.174** (0.009)	-0.053** (0.014)	-0.128** (0.009)	-0.387** (0.042)	-0.063** (0.017)	-0.026 (0.031)
$HSshare^{k}_{i,t-1}$	-0.871** (0.128)	-1.080** (0.236)	-0.605** (0.122)	-0.935** (0.345)	-1.153** (0.300)	-1.133** (0.307)	-0.444** (0.131)
Constant	-0.266** (0.011)	-0.200** (0.027)	-0.692** (0.032)	0.618** (0.056)	-0.447** (0.042)	-0.639** (0.059)	-0.745** (0.052)
Selection equation (exporting decision)							
$\ln GDPpc_{j,t-1}$	-0.217** (0.005)	-0.344** (0.008)	-0.015** (0.006)	-1.547** (0.025)	0.047** (0.012)	-0.149** (0.038)	-0.016 (0.015)
$pshare_{ij,t-1}$	-1.153** (0.028)	-1.386** (0.034)	-0.106 (0.063)	-1.360** (0.035)	0.891** (0.307)	-0.178** (0.068)	0.548** (0.196)
$HSshare^{k}_{i,t-1}$	-1.956** (0.783)	11.652* (6.131)	-3.312** (0.689)	22.625** (7.335)	-2.294 (2.604)	-6.558** (1.628)	-0.493 (2.106)
Constant	3.766** (0.047)	5.014** (0.077)	1.818** (0.056)	17.452** (0.261)	1.778** (0.092)	3.198** (0.394)	1.794** (0.116)
Uncensored observations	948891	691488	257403	385500	305988	149859	107544
Censored observations	52205	39211	12994	34347	4864	7755	5239
Wald Chi2-test of model validation	5845.68**	4986.25**	1505.54**	2677.13**	2160.87**	1049.49**	502.52**
Wald Chi2-test of independent equations	962.96**	799.46**	26.58**	373.90**	1.44	30.69**	0.21

Robust standard errors in parentheses. * significant at 5%; ** significant at 1%. † different from 1 at 5%.
The test of independent equations compares the joint likelihood of an independent probit model for the selection equation and a regression model on the observed price data against the Heckman model likelihood. A significant Chi2 value means that the residuals of the regression equation and of the selection equation are correlated, which justifies the use of the Heckman selection model.

Table 4: panel fixed effects with common markets							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	all	China	India	China to high_inc	China to low_inc	India to high_inc	India to low_inc
$\Delta \ln\left(e_{ij,t-1}\right)$	-1.191**† (0.030)	-1.158**† (0.034)	-0.587**† (0.122)	-1.090**† (0.044)	-1.216**† (0.054)	-0.446**† (0.160)	-0.804**† (0.189)
$\ln GDPpc_{i,t-1}$	0.043** (0.004)	0.012* (0.005)	0.121** (0.010)	-0.016* (0.007)	0.051** (0.008)	0.155** (0.015)	0.108** (0.015)
$\ln GDPpc_{j,t-1}$	0.064** (0.006)	0.078** (0.007)	-0.024 (0.013)	0.076** (0.013)	0.069** (0.009)	-0.101** (0.024)	0.013 (0.015)
$\text{var}\left[\Delta \ln\left(e_{ij,t-1}\right)\right]$	16.164** (1.478)	-0.779 (2.267)	33.387** (2.901)	-6.513* (3.031)	2.979 (3.459)	43.517** (3.771)	18.309** (4.556)
$pshare_{ij,t-1}$	-0.386** (0.023)	-0.403** (0.026)	-0.318** (0.052)	-0.333** (0.027)	-1.327** (0.115)	-0.325** (0.060)	-0.286** (0.107)
$HSshare^{k}_{i,t-1}$	-2.791** (0.448)	-5.833** (0.976)	-1.923** (0.511)	-6.360** (1.882)	-3.030** (1.027)	-2.686* (1.187)	-1.777** (0.567)
Constant	-0.837** (0.047)	-0.723** (0.061)	-0.547** (0.085)	-0.576** (0.126)	-0.852** (0.067)	0.044 (0.176)	-0.774** (0.094)
Observations	597507	421616	175891	204206	217410	97425	78466
Importer-product groups	97172	63267	33905	24193	39074	16590	17315
F-test	625.09**	499.65**	154.89**	252.50**	257.86**	111.11**	50.30**
Robust standard errors in parentheses. * significant at 5%; ** significant at 1%. † different from 1 at 5%.							

Table 5: panel fixed effects with asymmetry							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	all	China	India	China to high_inc	China to low_inc	India to high_inc	India to low_inc
$\Delta^{+} \ln\left(e_{ij,t-1}\right)$	-1.054**† (0.026)	-0.993** (0.029)	-0.611**† (0.122)	-0.972** (0.036)	-1.050** (0.049)	-0.441**† (0.157)	-0.875** (0.192)
$\Delta^{-} \ln\left(e_{ij,t-1}\right)$	-1.050** (0.026)	-0.989** (0.029)	-0.610**† (0.122)	-0.967** (0.036)	-1.045** (0.049)	-0.437**† (0.158)	-0.877** (0.193)
$\ln GDPpc_{i,t-1}$	0.059** (0.004)	0.036** (0.005)	0.115** (0.009)	0.034** (0.006)	0.050** (0.008)	0.127** (0.014)	0.122** (0.014)
$\ln GDPpc_{j,t-1}$	0.054** (0.006)	0.073** (0.007)	-0.010 (0.012)	0.025* (0.010)	0.095** (0.009)	-0.047* (0.021)	0.004 (0.014)
$\text{var}\left[\Delta \ln\left(e_{ij,t-1}\right)\right]$	15.092** (1.234)	2.106 (1.832)	30.282** (2.358)	-1.238 (2.311)	5.186 (3.009)	35.203** (2.968)	20.134** (3.891)
$pshare_{ij,t-1}$	-0.311** (0.020)	-0.423** (0.026)	-0.146** (0.032)	-0.337** (0.026)	-1.067** (0.100)	-0.152** (0.035)	-0.095 (0.079)
$HSshare^{k}_{i,t-1}$	-2.802** (0.364)	-3.727** (0.669)	-1.924** (0.415)	-4.107** (0.925)	-3.442** (0.872)	-3.757** (1.182)	-1.713** (0.459)
Constant	-0.866** (0.040)	-0.874** (0.052)	-0.634** (0.074)	-0.420** (0.088)	-1.064** (0.061)	-0.328* (0.149)	-0.797** (0.079)
Observations	948891	691488	257403	385500	305988	149859	107544
Importer-product groups	154253	107195	47058	55293	54327	23556	23502
F-test	862.77**	700.43**	196.24**	403.12**	289.99**	143.41**	59.34**
Robust standard errors in parentheses. * significant at 5%; ** significant at 1%. † different from 1 at 5%.							

Table 6: panel fixed effects with wages							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	all	China	India	China to high_inc	China to low_inc	India to high_inc	India to low_inc
$\Delta \ln\left(e_{ij,t-1}\right)$	-1.098** (0.046)	-0.987** (0.051)	-0.864** (0.166)	-1.000** (0.067)	-0.991** (0.076)	-0.841** (0.186)	-0.982** (0.367)
$\ln w_{i,t-1}^{k}$	0.029** (0.005)	0.046** (0.007)	-0.022* (0.010)	0.013 (0.009)	0.075** (0.011)	-0.008 (0.011)	-0.072** (0.023)
$\ln w_{j,t-1}^{k}$	0.030** (0.005)	0.016** (0.006)	0.058** (0.009)	0.007 (0.005)	0.029** (0.011)	0.063** (0.010)	0.045 (0.026)
$\text{var}\left[\Delta \ln\left(e_{ij,t-1}\right)\right]$	13.069** (2.122)	1.612 (3.272)	31.803** (4.480)	-1.839 (4.433)	4.025 (4.792)	29.168** (4.949)	39.332** (10.232)
$pshare_{ij,t-1}$	-0.327** (0.045)	-0.906** (0.120)	-0.210** (0.048)	-0.938** (0.135)	-0.878** (0.204)	-0.222** (0.050)	0.019 (0.158)
$HSshare_{i,t-1}^{k}$	-3.500** (0.772)	-4.543** (1.237)	-0.900 (0.550)	-4.672* (2.055)	-4.663** (1.463)	-3.476** (1.202)	-0.614 (0.643)
Constant	-0.000 (0.010)	0.038** (0.010)	-0.152** (0.031)	0.050** (0.017)	0.046** (0.007)	-0.169** (0.039)	-0.157** (0.046)
Observations	312602	223500	89102	100909	122591	67022	22080
Importer-product groups	63966	45461	18505	17305	29049	11948	6557
F-test	205.38**	184.23**	50.35**	86.91**	101.13**	46.81**	8.26**
Robust standard errors in parentheses. * significant at 5%; ** significant at 1%. † different from 1 at 5%.							

Appendix
Table A1: Export price data availability for high and low-income markets

High-income (1994-2007 GDP per capita average higher than 10,000USD)				Low-income (1994-2007 GDP per capita average lower than 10,000USD)			
China		India		China		India	
Australia	34760	Canada	22919	Argentina	23896	Argentina	8469
Austria	15292	Hong Kong	19429	Brazil	26461	Brazil	11508
Belgium	21433	France	25938	Bulgaria	12475	Chile	8547
Canada	36431	Germany	33225	Chile	25746	China	16211
Hong Kong	57011	Israel	15027	Colombia	17638	Colombia	6358
Cyprus	13147	Italy	27169	Czech Rep.	15717	Egypt	16198
Denmark	19485	Japan	23943	Egypt	27193	Indonesia	17997
Finland	18656	Korea Rep.	16585	Estonia	8347	Iran	11821
France	33845	USA	42905	Hungary	15739	Jordan	10553
Germany	40910	UK	37458	India	31805	Malaysia	25845
Greece	21764			Indonesia	38279	Mexico	10882
Iceland	4756			Iran	22991	Morocco	6836
Ireland	13541			Jordan	18453	Pakistan	9030
Israel	26480			Latvia	9909	Peru	4818
Italy	37066			Lithuania	10568	Philippines	12838
Japan	53661			Malaysia	35607	Russian Fed.	12413
Luxembourg	2753			Mexico	22063	South Africa	16719
Malta	9106			Morocco	16582	Thailand	20109
Netherlands	30015			Pakistan	26052	Tunisia	4224
New Zealand	24535			Peru	16974	Turkey	14756
Norway	16346			Philippines	30580	Viet Nam	10218
Portugal	16605			Poland	20321		
Korea Rep.	5420			Romania	16184		
Singapore	37356			Russian Fed.	28005		
Slovenia	10529			Slovakia	8097		
				South Africa	20990		

Table A2: Average export price data for high and low-income markets

High-income (1994-2007 GDP per capita average higher than 10,000USD)				Low-income (1994-2007 GDP per capita average lower than 10,000USD)			
China		India		China		India	
Australia	8415	Canada	410	Argentina	1750	Argentina	663
Austria	396	Hong Kong	1690	Brazil	4221	Brazil	863
Belgium	11740	France	1894	Bulgaria	637	Chile	1188
Canada	6061	Germany	1219	Chile	2143	China	3001
Hong Kong	5046	Israel	1277	Colombia	3564	Colombia	472
Cyprus	14584	Italy	746	Czech Rep.	3189	Egypt	4295
Denmark	19221	Japan	748	Egypt	3767	Indonesia	1298
Finland	5059	Korea Rep.	1030	Estonia	699	Iran	3038
France	9468	USA	2188	Hungary	396	Jordan	543
Germany	11896	UK	1557	India	3897	Malaysia	1315
Greece	18235			Indonesia	9841	Mexico	604
Iceland	3572			Iran	26962	Morocco	593
Ireland	3037			Jordan	1474	Pakistan	518
Israel	948			Latvia	724	Peru	397
Italy	7039			Lithuania	520	Philippines	630
Japan	6572			Malaysia	7633	Russian Fed.	1904
Luxembourg	923			Mexico	3042	South Africa	7954
Malta	43365			Morocco	1381	Thailand	1124
Netherlands	6647			Pakistan	5421	Tunisia	1070
New Zealand	814			Peru	1951	Turkey	5290
Norway	29019			Philippines	5745	Viet Nam	2942
Portugal	746			Poland	4586		
Korea Rep.	13202			Romania	1565		
Singapore	11442			Russian Fed.	4460		
Slovenia	462			Slovakia	3609		
				South Africa	2487		
Average	8580	Average	1355	Average	4709	Average	2199

Table A3: Average export price data for two-digit products

Products	China	India
01-05 Animal & Animal Products	28.77	281.83
28-38 Chemicals & Allied Industries	73.93	37.50
16-24 Foodstuffs	55.45	2.39
64-67 Footwear / Headgear	9.19	16.76
84-85 Machinery / Electrical	8116.39	3983.74
72-83 Metals	10.08	8.27
25-27 Mineral Products	1.56	4.67
39-40 Plastics / Rubbers	10.36	8.30
41-43 Raw Hides, Skins, Leather, & Furs	19.96	19.97
68-71 Stone / Glass	144.81	319.15
50-63 Textiles	9.99	9.88
86-89 Transportation	230700.80	42188.85
06-15 Vegetable Products	4.86	12.30
44-49 Wood & Wood Products	22.72	12.54