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Pricing-to-market, Trade Policy, and Market Power *

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Abstract. This paper studies the determinants of pricing-to-market at the firm-level, with a particular focus on the role of firm-specific and policy-induced market power. We use a large dataset containing export values and quantities by product and destination for all exporting firms in 12 developing and emerging countries, over several years. We first show that firms in our sample do price to market, i.e. significantly adjust their unit values in home currency in response to exchange-rate variations. The extent of pricing-to-market is quantitatively limited but highly significant and homogenous across origin countries despite their very different levels of development. We then study how firm performance and trade policy affect pricing-to-market at the firm-level. We find that within a given origin-destination-product cell, large, high-performance exporters price more to market. More importantly, we identify significant effects of trade-policy instruments on pricing-to-market: Higher import tariffs on a destination market are associated with less pricing-to-market, whereas non-tariff measures are associated with more. These results are consistent with models where pricing-to-market is increasing in firm size and market share, and suggest that trade policy has deep effects on market power, the direction of which depends on the type of instrument used.

JEL classification: F12, F13, F14, F31.
Keywords: Pricing-to-market, trade policy, exchange rate, tariffs

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1 Introduction

The idea that market power plays a strong role in determining pricing-to-market (PTM) behavior at the firm-level goes back at least to Krugman (1986). Yet, it is only recently that research has provided direct empirical evidence that large, productive firms with high market shares price more to market.[1] This research is however mostly limited to industrial countries and there is still much to understand about the way in which market power and market structure affect PTM at the firm-level.

We contribute to this literature in two main ways. First, we use the first large set of firm-level data covering several low- and middle-income countries and show that the literature's key results extend to a more general setting, even after controlling for a powerful array of fixed effects. Second and more importantly, we study how trade policy affects the extent of PTM. We can read our results backward: Assuming that differences in PTM reflect differences in market power, the response of individual firms to exchange-rate shocks can tell us something about competition and market structure in their destination markets, and by implication, about the effects of trade policies on market structure. We provide evidence that trade policy has indeed deep effects on firms' market power, and that those effects can go in opposite directions, depending on which policy instrument is used.

As a guide to our empirical analysis of the role of market power for PTM, we use a simple theoretical framework based on Atkeson and Burstein (2008) and derive testable predictions on the extent of PTM at the firm-level and how it relates to firm size and destination-market structure. Specifically, we show how the presence of tariffs in certain destination markets but not others generates PTM, although the effect is non-monotone in firm size. For firms with small market shares, tariffs unambiguously raise pass-through, while the effect becomes indeterminate as size/market share increases. As an extension, we also look at non-tariff measures, like technical or sanitary regulations. As tariffs act through rent-shifting effects we expect non-tariff measures applied in a non-discriminatory way – that is in compliance with the WTO's "national treatment" clauses, whereby imported and domestically-produced products must be treated alike – either to have no effect on PTM or to reduce it for incumbents if such measures induce the exit of smaller firms through higher fixed costs.[2]

We test these predictions using a very large, multi-country firm-level dataset obtained from customs administrations in twelve countries at different stages of development, ranging from low income in the case of Uganda to OECD in the case of Mexico. The ability to pool together firm-level data from several countries is a first and lends itself to a systematic exploration of the drivers of PTM.

Two main findings emerge. First, there is statistically significant but quantitatively limited PTM in all of the sample's regions of origin. Interestingly, the adjustment of origin-currency export prices to real exchange-rate changes turns out to be very similar across origin countries in spite of their very different sizes and income levels. Our estimates suggest that following a 10 percent exchange rate depreciation firms increase their home-currency price by 0.5 to 1.5 percent

[1] See for example Atkeson and Burstein (2008), Berman, Martin and Mayer (2012) or Amiti, Istkhoki and Konings (2014).

[2] We will leave aside the case of quantitative restrictions, as those have largely been phased out, and focus on regulations, either sanitary or technical, of which there is a plethora in high- and middle-income countries.

depending on the origin country. This might be an indication that the deep determinants of PTM at the firm-level are similar across countries. In particular, given the skewness of the size distribution of exporters (Mayer and Ottaviano, 2007, Freund and Pierola, 2012), if firms react heterogeneously to exchange rates, one would mechanically expect to find low PTM elasticities in firm-level estimations as small firms, which adjust less at the price margin, represent the majority of observations in all countries, driving down the estimates. PTM also increases with various indicators of firm performance and this remains the case even after controlling for all country-pair-product determinants of PTM (i.e. after including origin-destination-product-year fixed effects).

Second, turning to the effect of trade policy, we find strong support for the model's predictions using data on disaggregated bilateral applied tariffs and non-tariff measures. In our sample of developing-country exporters, which includes predominantly firms with small shares of their destination markets, we find that tariffs significantly reduce PTM. As predicted by theory, this effect is weaker for firms in the sample with larger market shares. As for non-tariff measures, like technical or sanitary regulations, we find that firms adjust less their home-currency prices (i.e. do more pass-through) following exchange rate movements in markets where NTMs are less stringent. These effects are robust to controlling for firm-level indicators of performance and for destination and product-specific determinants of PTM. Quantitatively, the estimated effect of trade policy on PTM is large. For instance, while firms raise their prices by around 2% following a 10% exchange rate depreciation in tariff-free markets, no significant PTM is detected when tariffs are above 20-25%. In the case of NTMs, moving from zero to a 10% ad-valorem equivalent increases PTM by half, from 9 to 14%.

Overall, these results support the idea that market power is an important driver of PTM. More specifically, both the firm-specific component of market power, related to firms' performance, and the market-specific one, related to market structure, affect PTM. In the last part of the paper, we use tariffs and NTMs as instruments for market shares and confirm this interpretation: trade policy has a significant effect on market shares, and high market shares amplify PTM. While the argument that trade policy affects competition is an old one, there is to this day little firm-level evidence of how it does so and, in particular, how the effects of different trade-policy instruments play out. Our results suggest that useful information can be generated in this regard from the analysis of exporter adjustment to exchange-rate fluctuations, an idea already present in early works such Aw (1993) or Goldberg and Knetter (1999) (see the survey by Goldberg and Knetter, 1997).

Our paper contributes to a vast literature on the determinants of exchange rate pass-through and pricing-to-market. In standard monopolistic-competition models, markups are constant and "mill pricing" applies, so there can be no PTM; however, several extensions of the standard model can generate PTM, essentially with two additional ingredients: variable markups and trade costs. Variable markups can be obtained through a number of ways.[3] Corsetti and Dedola (2005) show that the introduction of per-unit distribution costs in an otherwise standard monopolistic competition model is sufficient to generate PTM.[4] This is because distribution costs

[3]See Burstein and Gopinath (2014) for a more general discussion.

[4]Burstein, Neves and Rebelo (2003) and Campa and Goldberg (2010) also discuss the presence of distribution costs in local currency.

create a disconnection between producer and consumer prices, lowering the price elasticity of demand perceived by the firms, which generates PTM. Alternatively, non-CES preferences such as quadratic preferences *à la* Melitz and Ottaviano (2008) also allow for variable markups along a linear demand curve.

Our approach follows Atkeson and Burstein (2008) and subsequent papers by Amiti, Itskhoki and Konings (2014) or Auer and Schoenle (2013). Their model features a two-level demand structure characterized by a higher elasticity of substitution at the lower level, combined with oligopolistic competition (either Cournot or Bertrand with product differentiation) also at the lower level, i.e. within each sector. In this setting, the price elasticity of demand faced by each firm varies with its market share, making optimal markups variable. In turn, trade costs generate different market shares across destinations (and between exports and domestic sales) implying different degrees of markup adjustment in response to firm-level (symmetric) cost shocks.

All these models generate PTM, the extent of which depends on the firms' market share, itself a correlate of productivity and size. Some of the models yield an unambiguous relationship between size and pricing behaviour, where large firms face or perceive a lower elasticity of demand, which makes their markups more responsive to exchange rate movements. For instance, in the model with distribution costs, the reason is that for more efficient firms, the additive distribution cost creates a relatively larger wedge between producer and consumer prices. In other models like Atkeson and Burstein (2008), the relationship between size and PTM is more complex and non-monotone. This is because the impact of firm-level price changes on the sectoral price index, which depends on size, can play at cross-purposes with the direct effect of size on the perceived elasticity of demand. In addition, Amiti, Istkhoki and Konings (2014) show that large exporters are also large importers. In their extension of the Atkeson-Burstein model, as imported inputs dampen the effect of exchange-rate shocks on costs, real hedging provides an additional channel of influence on PTM.

The relationship between PTM and market power highlighted in the theory also appears as an empirical regularity. The empirical literature on the relationship between pass-through and firm size goes back at least to the work of Feenstra, Gagnon and Knetter (1996).[5] More recently, a number of studies have provided evidence for this link using firm-level data in high-income and emerging countries. Berman, Martin and Mayer (2012) and Amiti, Itskhoki and Konings (2014) use a combination of firm-level export and balance-sheet data, respectively for France and Belgium, and find that large, more efficient exporters do more PTM. Chatterjee, Dix-Carneiro and Vichyanon (2013) and Li, Ma and Xu (2013) provide similar evidence using respectively Brazilian and Chinese firm-level data. We confirm that the results established in those papers on individual countries hold on a larger and more diversified set of countries, some low, some middle-income.

The remainder of the paper is structured as follows. The next section presents a summary model drawing on Atkeson-Burstein (2008), Amiti, Itskhoki and Konings (2014), and Auer and Schoenle (2013) and derive three main predictions to guide the empirical analysis. Section 3 presents our multi-country firm-level dataset and the following three sections tests our predictions: Section 4 estimates the extent of PTM at the firm-level across the different countries of our sample; Section 5 estimates how this average level of PTM varies across firms; And section

[5] See also Alessandria (2004) and Garetto (2012).

6 examines the effect of trade policy on PTM. Each section discusses the empirical approach, presents the main results, and offers some robustness checks. Finally, section 7 concludes.

2 A model à la Atkeson-Burstein

In order to guide our empirical analysis we derive here basic expressions for pass-through and pricing-to-market from a variable-markup model à la Atkeson and Burstein (2008, henceforth AB). We follow their treatment (and that of Amiti, Itskhoki and Konings 2014, which is similar) very closely with one difference.As our dataset does not include firm-level cost data, unlike AB we use shocks on bilateral exchange rates as the identification device for PTM. In order to stay close to the empirics, the exposition in this section derives an exchange-rate pass-through parameter at the firm-level and shows how it depends on firm size and destination-market structure. In our setting, incomplete pass-through of exchange-rate shocks implies PTM as it drives wedges between export prices and domestic ones, and between export prices to different destinations; we will thus use the two terms interchangeably.

Let stars denote variables expressed in foreign currency and hats denote log-changes. Let μ_{fdpt} be firm f's markup over marginal cost when it sells product p to destination d at time t, c_{fpt} and $c^{\star}_{fpt}$ its marginal cost expressed in home and foreign currencies respectively, and e_{odt} the exchange rate between origin o (firm f's home country) and destination d, expressed as home currency per unit of the foreign currency.[6] Firm f's foreign-currency price is

$$p^{\star}_{fdpt} = \mu_{fdpt}\, c^{\star}_{fpt} = \mu_{fdpt}\, c_{fpt}/e_{odt} \tag{1}$$

Log-differentiating (1) with respect to an exchange-rate shock while the home-currency marginal cost c_{fpt} is held constant gives

$$\hat{p}^{\star}_{fdpt} = \hat{\mu}_{fdpt} - \hat{e}_{odt}. \tag{2}$$

That is, a depreciation in the home currency is like a negative shock on firm f's marginal cost expressed in foreign currency. Let $s^{\star}_{fdpt}$ be firm f's share of the market for product p in destination country d at time t, defined in foreign-currency terms; i.e.

$$s^{\star}_{fdpt} = \frac{p^{\star}_{fdpt} q_{fdpt}}{P^{\star}_{dpt} Q_{dpt}} \tag{3}$$

where $P^{\star}_{dpt}$ and Q_{dpt} are CES aggregators for prices and quantities respectively in the destination market, the former expressed in destination currency (hence the star), and the latter in quantity units. Let

$$\Gamma_{fdpt}(s) = -\left.\frac{d\ln(\mu_{fdpt})}{d\ln(s)}\right|_{P^{\star}_{dpt}\text{constant}} = \frac{s}{1-[(1-s)/\rho]-(s/\eta)}\left(\frac{1}{\rho}-\frac{1}{\eta}\right) \tag{4}$$

be the elasticity of firm f's markup to that market share in a quantity-setting game. Expression (4) is derived in the appendix. Note that Γ is increasing in s, that $\Gamma(0)=0$, and that $\Gamma(1)$ is

[6] We use the term 'product' in order to stay close to the empirics; but p corresponds to what the literature calls a 'sector' and an (f,k) couple is what it calls a 'variety'.

finite. Then (2) can be rewritten as

$$\hat{p}^{\star}_{fdpt} = \Gamma_{fdpt}\hat{s}^{\star}_{fdpt} - \hat{e}_{odt}; \tag{5}$$

while the log-change in firm f's market share can itself be expressed (see appendix) as a function of the log-change in its price relative to the log-change in destination d's sectoral price index:

$$\hat{s}^{\star}_{fdpt} = (1-\rho)\left(\hat{p}^{\star}_{fdpt} - \hat{P}^{\star}_{dpt}\right) \tag{6}$$

where ρ is the elasticity of substitution between firm-specific varieties of product p (the lower level in AB's nested CES structure). Combining (5) and (6), the change in firm f's foreign-currency price is

$$\hat{p}^{\star}_{fdpt} = \frac{1}{1+(\rho-1)\Gamma_{fdpt}}\left[(\rho-1)\Gamma_{fdpt}\hat{P}^{\star}_{dpt} - \hat{e}_{odt}\right]. \tag{7}$$

Let

$$\lambda_{fdpt} = \frac{(\rho-1)\Gamma_{fdpt}}{1+(\rho-1)\Gamma_{fdpt}} \tag{8}$$

As $\rho > 1$, λ_{fdpt} is an increasing function of Γ. Noting that $p_{fdpt} = e_{odt}\, p^{\star}_{fdpt}$, the change in firm f's home-currency price can be written, after some rearrangement, as

$$\hat{p}_{fdpt} = \hat{p}^{\star}_{fdpt} + \hat{e}_{odt} = \lambda_{fdpt}\left(\hat{e}_{odt} + \hat{P}^{\star}_{dpt}\right). \tag{9}$$

Consider the effect of an increase in the exporters' exchange rate, $\hat{e}_{odt} > 0$. Then $\hat{P}^{\star}_{dpt} < 0$ (destination d's sectoral price index, in foreign currency, goes down), so $\hat{e}_{odt}$ and $\hat{P}^{\star}_{dpt}$ have opposite signs. However, if home firms have less than a hundred-percent market share in destination d, $\mid \hat{P}^{\star}_{dpt} \mid < \hat{e}_{odt}$, so the term in parentheses is positive. Moreover, as $0 < \lambda_{fdpt} < 1$ whenever $s^{\star}_{fdpt} > 0$, (9) implies some pass-through but not complete (pass-through would be complete with $\lambda_{fdpt} = 0$). Letting $\beta_{fdpt} = d\ln(p_{fdpt})/d\ln(e_{odt})$, we have the immediate result that

Proposition 1 (Incomplete pass-through): $0 < \beta_{fdpt} < 1$ for all active exporters.

In Section 4, we will provide new evidence on the size of β from within firm-product-destination estimation on our sample of developing-country firms and compare it with existing estimates from industrial countries.

Consider now the effect of firm size on the pass-through coefficient. By (5), given that $\Gamma(0) = 0$,

$$\lim_{s^{\star}_{fdpt}\to 0} \hat{p}^{\star}_{fdpt} = -\hat{e}_{odt}, \tag{10}$$

so pass-through is complete for very small firms. Similarly, for a very large firm ($s_{fdpt} \to 1$), $\hat{p}^{\star}_{fdpt} \to \hat{P}^{\star}_{dpt}$ so, by (6), $\hat{s}^{\star}_{fdpt} \to 0$; as $\Gamma(1)$ is finite, (5) implies that

$$\lim_{s^{\star}_{fdpt}\to 1} \hat{p}^{\star}_{fdpt} = -\hat{e}_{odt}. \tag{11}$$

Thus, both very small and very large firms tend to full pass-through, which is non-monotone as a function of firm size/market share, as discussed in detail by Auer and Schoenle (2013). Can we say anything more to guide the firm-level empirics? Consider the following thought experiment. In case 1, we consider the adjustment of a firm to a given positive exchange-rate shock. In case 2, we suppose that before the exchange-rate shock, the firm had a shock raising its marginal cost, whereas all other firms are as before. This is equivalent to considering two firms selling in the same destination-product-year (*dpt*) cell and differing only in their marginal cost, firm one being larger. *Ceteris paribus*, differences in marginal cost map one-to-one into differences in market share, so $s_{1dpt} > s_{2dpt}$. By abuse of notation, let $\hat{P}^{\star}_{idpt}$, $i = 1, 2$, be the adjustment in $\hat{P}^{\star}_{dpt}$ triggered by the combined effect of the exchange-rate change in case 1 and in case 2. Let also $\lambda_i = \lambda_{fdpt}$ in case i. Then

$$\begin{aligned} \hat{p}_{1dpt} - \hat{p}_{2dpt} &= (\lambda_1 - \lambda_2)\hat{e}_{odt} + \lambda_1 \hat{P}^{\star}_{1dpt} - \lambda_2 \hat{P}^{\star}_{2dpt} \\ &= (\lambda_1 - \lambda_2)(\hat{e}_{odt} + \hat{P}^{\star}_{1dpt}) + \lambda_2(\hat{P}^{\star}_{1dpt} - \hat{P}^{\star}_{2dpt}). \end{aligned} \tag{12}$$

If firm 1 is larger, by (4) and (8), $\lambda_1 - \lambda_2 > 0$. Again, $\hat{P}_{1dpt}$ and $\hat{e}_{ot}$ have opposite signs, but $\hat{e}_{odt} + \hat{P}_{idpt} > 0$, so the whole first term on the RHS of (12) is positive, contributing to a stronger adjustment of the larger firm's home-currency price (less pass-through). By contrast, the second term is negative, as $\lambda_2 > 0$ and $\hat{P}^{\star}_{1dpt} < \hat{P}^{\star}_{2dpt} < 0$ if firm 1 is larger than firm 2. Thus, the general direction of the effect is indeterminate, but if firm 2 is small enough, the first term dominates and the larger one does less pass-through. Note also that the effect is unambiguous if sectoral price indices are held constant. Thus, we can state

Proposition 2 (Heterogeneous PTM): In general, the effect of size (market share) on firm-level PTM is indeterminate. However, among small firms, the relatively larger ones do more PTM.

In the empirics, we will verify monotone sorting in terms of size in the case of a fixed sectoral price level by estimating pass-through within product-destination-year cells, and in the general case (without a priori) by estimating it within firm-product-destination cells.

Consider finally two destination markets that are identical in every respect except that firm f faces a tariff in market 2 and none in market 1. Let us also assume that the exchange-rate shock is on the exporter's home currency, entailing a symmetric devaluation in terms of all destination currencies; that is, $\hat{e}_{o1t} = \hat{e}_{o2t} = \hat{e}_{ot}$. Then we have again

$$\hat{p}_{f1pt} - \hat{p}_{f2pt} = (\lambda_1 - \lambda_2)(\hat{e}_{ot} + \hat{P}_{1pt}) + \lambda_2(\hat{P}_{1pt} - \hat{P}_{2pt}). \tag{13}$$

Provided that the RHS of (13) is nonzero, there is pricing-to-market in the sense that firm f adjusts the home-currency price of product p differently across destinations in response to a symmetric shock. Under the assumption that destination markets are symmetric up to the tariff, firm f's market share is higher in market 1, where it faces no tariff, than in 2, where it faces one; so $\lambda_1 > \lambda_2$ again. This direct effect of the tariff on firm f's pricing contributes to *less* pass-through on market 1 than on market 2.[7]

[7] This effect implies that pass-through is higher in the presence of a tariff. AB note that "[t]he force of this effect in our model is in the wrong direction—through this effect a cost shock to a home firm leads it to raise its export price relative to its domestic price." (p. 2024). In our setting, the tariff-ridden market 2 is like AB's export

As before, the indirect effect generated by strategic interaction between firms, picked up by the second term on the RHS of (13), is indeterminate. On one hand, the direct effect means that firms adjust their foreign-currency price by less in market 1, which contributes to cut the sectoral price index $P^{\star}_{dpt}$ by less in market 1; that is, $\hat{P}^{\star}_{2pt} < \hat{P}^{\star}_{1pt} < 0$, reinforcing the direct effect. On the other hand, home firms subject to the shock have a higher share of market 1, which, for given pass-through, contributes to cut $P^{\star}_{dpt}$ by more in market 1. This, by contrast, works at cross-purposes with the direct effect. Which effect dominates on firm f's pricing depends on its market share and on the combined market share of origin-o exporters, all subject to the same exchange-rate shock, in destination d. If firm f's market share is sufficiently small, its price adjustment has no impact on the destination's sectoral price index, leaving only the direct effect $(\lambda_1 - \lambda_2)\hat{e}_{ot}$, in which case the tariff raises the degree of pass-through. Thus, we have

Proposition 3 (Tariffs and pricing-to-market): Tariffs induce PTM, i.e. heterogeneous price adjustment to common exchange-rate shocks across symmetric destinations. The direction of the effect is, in general, indeterminate; however, for small firms, higher tariffs, ceteris paribus, induce less PTM.

Given that our dataset consists of a majority of relatively small developing-country exporters, that is what we expect to find in the data. Finally, given that the result in Proposition 3 is based on induced changes in market shares, i.e. on the rent-shifting effect of tariffs, we can state:

Corollary 1 (Non-tariff measures and pricing-to-market): Non-tariff measures (technical and sanitary regulations) applied in a non-discriminatory way either have no effect on pass-through or reduce it for incumbents if they induce the exit of smaller firms.

Whether non-tariff measures are discriminatory or not is an important issue in trade-policy discussions. Corollary 1 shows that the direction of their effect on pass-through can be read as an indication of whether they have rent-shifting effects or not, i.e. whether they are applied in a discriminatory way or not. We will explore the issue in the last part of the empirics, using variation in their stringency from a new dataset on NTM ad-valorem equivalents.

3 Data and empirical methodology

For our analysis we combine three main types of data: (i) firm-level data on export flows, (ii) macroeconomic data and (iii) trade policy variables.

3.1 Firm-level trade data

Our data was obtained from the customs administrations of twelve developing countries. Data for Kenya, Rwanda, Tanzania and Uganda was obtained by the International Growth Center

market, while market 1 is like the home market. In our setting, faced with a shock equivalent to a marginal-cost reduction, the firm cuts its price more in the "foreign" (tariff-ridden) market than in the "home" (tariff-free) one. The logic is the same.

and data for Bangladesh, Chile, Jordan, Kuwait, Lebanon, Mexico, Morocco, and Yemen was obtained by the Trade and Integration Unit of the World Bank Research Department, as part of the Exporters Dynamics Database (EDD) project described in Cebeci et al. (2012). For each country, all export transactions are covered over a certain time-period (see Table 1). For each firm and year, the data includes a firm identifier, as well as the value (in local currency) and quantity (expressed in kilograms) sold by the firm for each destination country and HS product (at country-specific HS8-equivalent levels).[8] For each firm-destination-product-year, unit values are computed as the ratio of export value to quantity. We clean the data in a number of ways. First, we keep only flows over a thousand USD. Second, for both unit values and export volumes we drop all observations belonging to the top and bottom percentiles in terms of levels and growth rate, percentiles being computed by origin country and sector (HS2).

Table 1 gives basic information on our final sample size and sample periods, by origin country. The sample is dominated by four countries, Bangladesh, Chile, Mexico and Morocco, in terms of transactions (both total and yearly) and number of firms. All origin countries have diversified destination portfolios, and the total number of HS6 products exported in one year or another ranges between 126 (Rwanda) and 5,607 (Mexico) out of a notional total of about nearly 6,000 HS6 lines. Sub-Saharan African firms are less diversified on average in terms of both number of destinations and products. Differences in terms of diversification are particularly important in terms of number of products (total or averaged by firm).

Table 1: Sample characteristics

Country	Period	# observations	Obs./year	# firms	# dest.	# products	dest./firm	prod./firm
Bangladesh	2006-11	128,600	21,915	7,487	159	1,030	10.7	8.1
Chile	2004-09	205,842	34,675	6,526	158	3,120	15.9	9.8
Jordan	2004-11	22,490	3,086	2,074	141	1,142	9.9	5.1
Kenya	2006-11	42,759	7,802	2,921	139	2,296	9.4	15.2
Kuwait	2009-10	4,600	2,310	814	73	941	7.5	22.9
Lebanon	2009-10	30,261	15,134	2,508	133	1,692	11.9	33.3
Mexico	2001-09	587,897	86,973	48,621	160	5,607	8.9	23.9
Morocco	2003-10	125,371	15,702	6,295	154	2,374	8.4	12.4
Rwanda	2006-11	763	147	229	41	126	6.1	3.2
Tanzania	2006-11	7,089	1,335	987	100	775	7.9	6.6
Uganda	2005-11	6,294	1,005	709	81	635	7.7	6.4
Yemen	2007-10	2,180	735	425	59	285	8.9	15.7

Product are defined at the six-digit level of the Harmonized System.

3.2 Country-level variables

Exchange rates vis-a-vis the U.S. dollar are from the IMF's International Financial Statistics (IFS) and are deflated by consumer price indices to obtain real exchange rates (RER). They are all expressed in local currency units (LCU) per dollar in the IFS. Let e_{ot} and e_{dt} be respectively

[8] Product classifications are not harmonized between countries at sub-HS6 levels of disaggregation (HS8 or HS10). This is not a problem in our estimations as all regressions have fixed effects at the firm-destination-product level. However, for comparability of descriptive statistics, we aggregate products up to the harmonized HS6 level.

the origin and destination countries' exchange rates in LCU per dollar in year t, and p_{ot} and p_{dt} their consumer price indices. Our bilateral exchange-rate variable, in logs, is thus:[9]

$$\ln(e_{odt}) = \ln\left(\frac{e_{ot}/p_{ot}}{e_{dt}/p_{dt}}\right) = \ln\left(\frac{e_{ot}}{e_{dt}}\right) - \ln\left(\frac{p_{ot}}{p_{dt}}\right) \tag{14}$$

Finally, GDP data are from the World Bank's World Development Indicators (WDI).

3.3 Trade policy variables

We use data on both tariffs and non-tariff measures. For tariffs, we use data on Most Favored Nation (MFN) and preferential tariffs at the HS6 level from TRAINS. For each country-pair-product and year we compute the bilateral applied tariff and then take the average over the period in order to smooth out missing values (and given that we are interested in the role played by differences trade policy across markets, rather than in the effect of variations in trade policy in a given market over time).

For non-tariff measures we use ad-valorem equivalents (AVEs) at the destination country-product level estimated in Cadot and Gourdon (2014a, 2014b) to which we refer the reader for details. The source data was collected as part of a joint project of UNCTAD and the World Bank. It currently covers 45 countries and consists of binary indicators taking value one when measure of type j is applied to product p (defined at the six-digit level of the Harmonized System) by destination (importing) country d, and zero otherwise. Measures are coded according to the MAST (Multi-Agency Support Team) classification revised in 2012. The data covers sanitary and phytosanitary measures (SPS), technical barriers to trade (TBT) and other measures.[10] The binary data was converted into estimated ad-valorem equivalents (AVEs) by running OLS regressions of the log of trade unit values on NTM dummies (the family of binary indicators marking the application of each NTM type to each product by each destination country) and control variables (including bilateral distance, income levels, etc...). Country-specific estimates were obtained by interacting NTM dummies, by type of instrument, with importer dummies, allowing for different modalities of application of the same type of measure between different importing countries.

The bulk of the variations in NTM AVEs (more than 80%) is attributable to the application of SPS and TBT regulations. Table 9 in the appendix shows the (unweighted) average levels of MFN tariffs and NTMs AVEs for the 45 countries covered by both tariff and NTM data. The lowest levels of MFN tariffs are observed in developed countries (e.g. 3.1% in Japan), while North African and South Asian countries have the highest (over 20% in India, Tunisia or Morocco). NTM AVEs also differ across countries, China having the highest (25%), which seems to accord with anecdotal evidence of trade-restrictive application of regulatory measures. Countries with similar regulations (e.g. members of the European Union) may nevertheless have different AVEs if they enforce them differently, which is the case for some of the Eastern European members

[9]In our baseline estimations, we have dropped the top percentile of country-pairs in terms of variance of bilateral real exchange rates. Dropping these countries which display extreme price variations (generally countries with hyperinflation) limits measurement error and only drops 0.05 percent of total trade value.

[10]These include for e.g. trade-related investment measures or intellectual property, although data on those is very scant. For more information on the MAST nomenclature, see: http://unctad.org/en/PublicationsLibrary/ditctab20122_en.pdf.

(e.g. Hungary vs. the Czech Republic). Note that while some countries are characterized by high levels of both NTMs and tariffs, the overall correlation between the two is not statistically significant at common confidence levels.

3.4 Descriptive statistics

Table 2 shows descriptive statistics for the variables used in the regressions. Our final sample contains around 82,000 firms. Unsurprisingly, trade-policy variables have the largest proportion of missing values. Both tariffs and the estimated ad valorem equivalents of NTM are low around 5% to 7% on average.[11] The median firm in our sample exports 4 products and serves 2 destinations against 10 products and 2 destinations for the average firm. This skewness in the distribution of products and destinations is consistent with stylized facts documented by the literature over the last decade: most exporters export only one product to a single destination and exports are dominated by a few very large, multi-products, multi-destinations firms (see for instance Mayer and Ottaviano, 2007).[12]

Table 2: Descriptive statistics

	Obs.	Mean	S.D.	Q1	Median	Q3
Volume (weigth in kg)	1,164,146	1.72E+05	2.86E+06	6.25E+02	4.94E+03	3.80E+04
ln volume	1,164,146	8.41	3.00	6.44	8.51	10.54
Unit value (LCU)	1,164,146	1.01E+05	2.41E+07	4.55E+01	2.82E+02	1.31E+03
ln unit value	1,164,146	5.58	2.52	3.82	5.64	7.18
Number of products (firm, t0)	1,164,146	90.77	263.53	5.00	14.00	40.00
ln number of products	1,164,146	2.78	1.67	1.61	2.64	3.69
Number of destinations (firm, t0)	1,164,146	18.13	31.94	2.00	7.00	19.00
ln number of destinations	1,164,146	1.92	1.41	0.69	1.95	2.94
Real exchange rate	1,133,033	104.78	398.67	4.25	10.90	34.70
ln real exchange rate	1,133,033	2.09	2.78	1.45	2.39	3.55
GDP constant (USD 2000)	1,151,902	4.44E+12	5.44E+12	8.55E+10	1.17E+12	1.18E+13
ln GDP	1,151,902	27.27	2.61	25.17	27.78	30.10
Bilateral distance	1,164,146	4.79E+03	3.85E+03	2.06E+03	3.37E+03	6.62E+03
ln distance	1,164,146	8.17	0.81	7.63	8.12	8.80
Foreign import tariff	753,931	5.13	13.73	0.00	0.25	6.00
ln (tariff/100+1)	753,931	0.05	0.08	0.00	0.00	0.06
Non-tariff measure (NTM AVE)	205,462	0.07	0.29	0.00	0.00	0.07

The number of products and destinations are computed for each firm in the first year it enters the dataset. GDP data is reported for the destination country. Foreign import tariffs are computed as the average over the period of the corresponding country-pair-product-year applied tariffs in order to smooth out missing values. For non-tariff measures we use ad-valorem equivalents (AVEs) at the destination country-product level estimated from Cadot and Gourdon (2014a, 2014b). Products are defined at the six-digit level of the Harmonized System.

We now turn to an empirical exploration of the three predictions derived from our theoretical framework in section 2. For each prediction we present in a separate section the empirical approach, the main results and a series of robustness checks.

[11] These numbers are lower than the average levels of protection displayed in Table 9 in the appendix, which was expected as high levels of protection deter trade and are therefore less likely to be observed in our final dataset.

[12] In Table 2, the median numbers of (HS6) products and destinations appear respectively as 14 and 7. This reflects multiple counting of multi-product multi-destinations exporters at the level of the unit of observation (firm-destination-product).

4 Pricing-to-market across countries

In this section we start with proposition 1; we estimate the extent of PTM at the firm-level across the twelve origin countries in our sample and compare our estimates with results from the existing literature.

4.1 Econometric specification

Let us define the following indices: o is origin country, d is destination country, f is firm, p is product, and t is year. Let e_{odt} be the average real exchange rate between the origin and destination countries in year t as defined in (14) and UV_{fdpt} be the producer price of product p exported by firm f to destination d at t, in country o's currency (proxied by unit value). Note that as our dataset does not contain information on firms' ownership all firms in our sample are treated as independent entities. Thus in the presence of firm subscripts we omit the origin country subscripts. Let $\mathbf{x_{dt}}$ be a set of time-varying destination specific controls, including the destination's GDP. Finally, let $\mathbf{FE}_{ot}$ and $\mathbf{FE}_{fdp}$ be respectively origin-year and firm-destination-product fixed effects. The inclusion of these fixed effects implies in particular that we estimate the effect of the real exchange rate variations on firm-level prices within a given firm-destination-product triplet over time. Our baseline estimating equation for proposition 1 is:

$$\ln \text{UV}_{fdpt} = \beta_1 \ln e_{odt} + \gamma \mathbf{x_{dt}} + \mathbf{FE}_{ot} + \mathbf{FE}_{fpd} + \varepsilon_{fdpt}. \tag{15}$$

In equation (15) an increase in e_{odt} is a depreciation of the home currency of firm f (the exporter). In the presence of PTM, firm f will then raise its home-currency producer price and capture some of the rent generated by the depreciation. Thus, in the presence of PTM we expect the parameter estimate for β_1 to be positive and less than one.

Equation (15) is estimated by OLS with robust standard errors clustered at the dyad-product-year level.[13] Note that β may suffer from attenuation bias due to classical measurement error: because we use annual data, the exchange rates applied in our estimations are (potentially) not exactly the ones actually faced by the firms at the time they export. The only way to solve this issue would be to use higher frequency data which would allow us to match transactions with daily or monthly exchange rates. Indeed, Fosse (2012) shows using Danish data that moving from annual to monthly data increases the elasticity of unit values to the exchange rate (from 0.14 to 0.19 in his case).[14] This might explain why the literature using yearly trade customs data typically finds lower estimates of PTM than those found in papers using direct price data at a higher frequency (e.g. Gopinath and Itskhoki, 2010).

4.2 Results

Table 3 presents the results of our baseline regression for the whole sample (column 1) and split by country or country groups (columns 2 to 7). The dependent variable is the log of export unit values and all regressions are estimated by OLS with firm-product-destination and origin-year fixed effects. The PTM coefficient is the coefficient on the log of the real bilateral exchange rate.

[13]Our main results are robust to clustering the standard errors at the firm or at the dyad-year levels.
[14]See also Mallick and Marques (2010).

Two results are worth highlighting. First, on the whole sample (column 1) the elasticity of the exporter price to the log of the bilateral RER is positive and significant and implies that the average firm in our sample raises its price by 1.3% following a 10% depreciation of its home currency. This is the empirical counterpart of proposition 1. This elasticity is quantitatively close to the estimates reported in the literature on industrial and emerging countries, which typically lie between 0.05 and 0.2.[15]

Thus, although our elasticities are estimated on a sample of developing countries, we find a degree of PTM that is very consistent with those found for industrial countries. Moreover, like those found in the existing literature, our estimates also reflect limited PTM on average and very high levels of pass-through into export price (87% in column 1). The estimated PTM elasticities might be low due to measurement error arising from using annual data. An alternative explanation is that these firm-level estimates hide a great deal of heterogeneity. If large firms adjust more to exchange rate variations at the price margin PTM may be much stronger on aggregate. We test for the relationship between PTM and firms heterogenity in the next section.

Table 3: Exchange rates and unit values

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dep. var: ln(unit value)							
Exporting countries	All	Excl. Mexico	EAC	MENA	Bangladesh	Chile	Mexico
ln(RER)	0.129^{a}	0.119^{a}	0.124^{c}	0.051^{c}	0.126^{a}	0.129^{a}	0.143^{a}
	(0.011)	(0.012)	(0.067)	(0.027)	(0.028)	(0.014)	(0.021)
ln(dest. GDP)	0.054^{a}	-0.011	-0.193^{b}	-0.061	0.021	0.064^{b}	0.420^{a}
	(0.020)	(0.022)	(0.085)	(0.039)	(0.059)	(0.028)	(0.056)
Observations	1123384	562255	54684	182035	127600	197936	561129
Adj. R^2	0.042	0.090	0.088	0.072	0.216	0.054	0.006

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. All estimations include origin-year dummies and firm-destination-product fixed effects.

The second and perhaps more surprising result is that the degree of PTM is very homogenous across origin countries. In all cases the elasticity of producer prices to the exchange rate lies between 0.05 and 0.15. This might be an indication that after controlling for time-invariant firm-product-destination characteristics and supply-side shocks, the deep determinants of PTM at the firm-level are similar across countries. In particular, and related to the previous point, given that exports are systematically concentrated among a few very large firms, if firms react heterogeneously to exchange rates, one would mechanically expect to find low PTM elasticities in firm-level estimations as small firms, which adjust less at the price margin, represent the majority of observations, driving down the estimates.

[15] Around 0.1 in France (Berman et al., 2012); 0.15 in Denmark (Fosse, 2012); 0.2 in Belgium (Amiti et al., 2014) and Brazil (Chatterjee et al., 2013); 0.06 in China (Li et al., 2013); 0.05 in Italy (Bernard et al., 2013).

4.3 Pricing-to-market across destinations

Does PTM differ across destination countries? If competition is tougher in high-income countries (and therefore individual market shares are smaller) we expect to observe less PTM–this would be the prediction of Atkeson and Burstein (2008), but also of models featuring linear demand such as Melitz and Ottaviano (2008). Alternatively, high-income countries might be characterized by larger distribution margins which have been shown to raise the incentive for PTM (Campa and Golberg, 2010; Corsetti and Dedola, 2005). We would also expect more PTM in high-income countries if PTM was increasing with quality and high-income markets demanded higher-quality products (Chen and Juvenal, 2014). Given these opposite forces, whether there should be more or less PTM in high-income countries is ultimately an empirical question, which we now test.

In table 4 we interact the exchange-rate variable with dummies marking destination-country income group: low income, lower-middle income, and upper-middle income, with high income being the omitted category. The distribution of firm sizes may differ across destination markets through selection effects. While much of that heterogeneity is controlled for by fdp (firm-destination-product) fixed effects, we exploit the high dimensionality of our dataset and further control for the heterogeneity in price adjustments across firms, destinations and products. Specifically we include interaction terms between the exchange rate and (i) exporter size (columns 2 to 4), (ii) origin-country dummies (columns 3 and 4), and (iii) product dummies (column 4). This ensures that the coefficient estimates on our interaction term between the exchange rate variable and the income group dummies indeed capture destination-specific effects.

Table 4: Pricing-to-market across destinations

Dep. var: ln unit value	(1)	(2)	(3)	(4)
ln(RER)	0.218^a	0.117^a		
	(0.017)	(0.023)		
ln(RER) × Upper middle income dest.	-0.102^a	-0.111^a	-0.127^a	-0.138^a
	(0.022)	(0.022)	(0.023)	(0.024)
ln(RER) × Lower middle income dest.	-0.311^a	-0.314^a	-0.324^a	-0.330^a
	(0.036)	(0.036)	(0.038)	(0.038)
ln(RER) × Low income dest.	-0.101	-0.109	-0.095	-0.111
	(0.067)	(0.066)	(0.077)	(0.075)
ln(dest. GDP)	0.053^a	0.060^a	0.062^a	0.060^a
	(0.020)	(0.020)	(0.020)	(0.020)
Additional interactions				
ln RER × firm size	No	Yes	Yes	Yes
ln RER × origin	No	No	Yes	Yes
ln RER × HS2	No	No	No	Yes
Observations	1123315	1123315	1123315	1123315
Adj. R^2	0.042	0.042	0.042	0.044

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by origin-destination-product-year are in parentheses. All estimations include origin-year dummies and firm-destination-product fixed effects.

The results are stable across specifications. PTM is non-monotonically related to income with the lowest level of PTM found in middle-income destinations. The high level of PTM found on high-income destination countries is consistent with the existing literature (e.g. Gaulier, Larheche-Revil and Mejean, 2008) and suggests that on such markets the forces of high distribution margins and/or higher-quality products seem to dominate the force of (presumably) stronger competition.

4.4 Robustness Checks

Table 10 in appendix checks the sensitivity of results in Table 3 replacing in equation (15) the *ot* (origin-year) fixed effects with *opt* (origin-product-year) ones. This allows to further control for origin-product time-varying effects such as comparative advantage or sectoral concentration. Results are robust to including origin-product-year effects.

We also estimate the equivalent of (15) on export volumes, i.e. replacing the left-hand side variable by the log of the quantity exported by firm f to country d in product p in year t. If PTM is less than complete, part of the exchange-rate variations are passed on to consumer prices in the destination countries. Thus, we also expect the coefficient on $\ln e_{odt}$ to be positive, as a depreciation of the real exchange rate with less than complete PTM should raise demand in the destination market and firm-level export volumes. Tables 11 and 12 in appendix replicate table 3 and table 10 using export volumes as the dependent variable. We find positive and significant but quantitatively limited elasticities to the exchange rate. Differences across origin countries are slightly larger, but the coefficients are also less precisely estimated.

5 Firm heterogeneity and pricing-to-market

In this section we turn to proposition 2 and estimate the relationship between PTM and firm-level heterogeneity.

5.1 Econometric specification

Our second prediction is that firm performance positively affects the extent of PTM. In Akteson and Burstein (2008) as shown in section 2, this is because high performance firms have larger market shares. This prediction more generally arises within a class of models featuring firm heterogeneity and variable markups. Large, high performance firms face or perceive a lower elasticity of demand which makes their markups more responsive to exchange rate movements.

Unfortunately, while our dataset is very large it is relatively poor in covariates as it contains no firm characteristics such as employment or value added. Thus, we rely on proxies for the identification of the effect of firm productivity or size.

In the literature product scope is the firm-level observable that correlates most closely across firms with productivity. However, within firms both the theoretical (see Bernard, Redding and Schott, 2011, Eckel and Neary 2010, Mayer, Melitz and Ottaviano, 2014) and the empirical literature (Chatterjee, Dix-Carneiro and Vichyanon, 2013) suggest that product scope is endogenous to the firm's environment. For instance, Bernard, Redding and Schott (2006) and Eckel and Neary (2010) show that firms optimally reduce product scope (focus on their core competencies) after trade liberalization as a result of pro-competitive effects. The same pro-competitive effects can be expected from an appreciation of the exporter's currency. Other proxies for firm performance such as total exports or the number of destinations served are even more clearly endogenous to exchange-rate variations. Thus, we have a problem of collinearity between firm-size proxies and exchange rates both on the right-hand side. Moreover, variations in firm performance and prices might be simultaneously affected by omitted variables.

We treat these problems in two ways. First, we measure our size proxies, product scope, destinations served and total export values, at the broader firm-level rather than at the firm-destination-product at which the regression is run. Second and more importantly, we use for each of those variables beginning-of-period values.

Letting φ_{f0} be firm f's performance at t_{f0}, the first year it enters the dataset, the corresponding estimating equation for proposition 2 is:

$$\ln \mathrm{UV}_{fdpt} = \beta_1 \ln e_{odt} + \beta_2 (\ln e_{odt} \times \ln \varphi_{f0}) + \gamma \mathbf{x_{dt}} + \mathbf{FE}_{ot} + \mathbf{FE}_{fdp} + \varepsilon_{fdpt} \tag{16}$$

where β_1 measures the average exchange-rate elasticity of unit values and β_2, the coefficient on the interaction term, measures the heterogeneity of reactions to exchange-rate variations between firms at different performance levels (the non-interacted term $\ln \varphi_{f0}$ is absorbed by the fixed effects $\mathbf{FE}_{fdp}$). The estimate for β_2 is expected to be positive if high-performance firms price more to market.

5.2 Results

Table 5 columns (1) to (3) report the OLS estimation results for equation (16) and can be thought of as the empirical tests of Proposition 2. In each column we use a different proxy for φ_{f0}, the number of products, the number of destinations, and the total export value, all taken at the beginning of each firm's sample period. Consistent with Berman *et al.* (2012) and Amiti *et al.* (2014) the results clearly suggest that large exporters react more in terms of prices than small ones. This result is robust to changes in the proxy of firm size used. Quantitatively, the heterogeneity in adjustment is non-negligible. The estimated coefficients in column 1 suggest that following a 10% depreciation of its home currency a firm exporting only one product will raise its price on average by only 0.3%, while a firm selling ten products will raise its price by 1.2% – three times more.

A potential issue with equation (16) is that the heterogeneity in PTM picked up by the coefficient β_2 could be driven by differences between products or destinations reflecting unobserved product or destination characteristics correlated with exporter size through selection effects. It might be the case, for instance, that high-performance firms export on average to more remote markets or to markets with higher distribution costs (higher distribution costs reduce the price elasticity of demand perceived by the firm and therefore encourage pricing-to-market).

To ensure that we indeed capture firm-performance effects and given the dataset high dimensionality we can go further than the existing literature and include in equation (16) origin-destination-product-year fixed effects. In that case, the exchange rate variable is absorbed by the fixed effects and the estimating equation becomes:

$$\ln \mathrm{UV}_{fdpt} = \beta_2 \left(\ln e_{odt} \times \ln \varphi_{f0}\right) + \gamma \mathbf{x_{dt}} + \mathbf{FE}_{odpt} + \mathbf{FE}_{fdp} + \varepsilon_{fdpt} \tag{17}$$

In this extremely demanding specification, β_2 unambiguously identifies the effect of firm characteristics on PTM, as all heterogeneity in PTM across products or destinations is controlled for by $\mathbf{FE}_{odpt}$. In other words, our interaction term now captures differences in PTM between firms of different sizes but located in the same origin country and selling the same product to the same destination in the same year. Another advantage of specification (17) is that is allows us to hold the price index in the destination country constant. As shown formally in section 2 this means that we focus only on the direct effect of firm size or firm market share on PTM filtering out the indirect effect of market share on PTM through adjustments of the price index. The drawback of this specification is that as the main variable e_{odt} is now absorbed by the fixed effects $\mathbf{FE}_{odpt}$, it is no longer possible to identify separately the average exchange-rate elasticity of unit values (our basic PTM elasticity).

Table 5 columns (4) to (6) show that the results are very robust – larger exporters do more pricing-to-market; if anything, they are reinforced quantitatively.[16]

Table 5: Firm heterogeneity and pricing-to-market

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. var.			ln unit value			
ln(RER)	0.032^c	0.033	-0.186^a			
	(0.018)	(0.027)	(0.071)			
ln(dest. GDP)	0.061^a	0.057^a	0.056^a			
	(0.020)	(0.020)	(0.020)			
ln(RER) × ln(# prod$_{t0}$)	0.036^a			0.063^a		
	(0.007)			(0.014)		
ln(RER) × ln(# dest$_{t0}$)		0.035^a			0.096^a	
		(0.009)			(0.018)	
ln(RER) × ln(export val$_{t0}$)			0.019^a			0.037^a
			(0.004)			(0.009)
Observations	1123290	1123290	1123289	1132939	1132939	1132938
Firm-product-destination FE	Yes	Yes	Yes	Yes	Yes	Yes
Origin-year FE	Yes	Yes	Yes	No	No	No
Origin-destination-product-year FE	No	No	No	Yes	Yes	Yes

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. ln(# prod$_{t0}$), ln(# dest$_{t0}$) and ln(export val$_{t0}$) are the number of products, the number of destinations and the total export value of the firm during the first year it appears in the sample.

[16] As an alternatively, we included a set of interaction terms between the exchange rate variable and destination dummies, and between the exchange rate and product dummies. Again, results were unchanged.

5.3 Robustness Checks

Table 13 in the appendix replicates table 5 taking export volumes as the dependent variable instead of unit values. The results clearly show that the exported volumes of high performance firms react less to exchange rate variations. In line with existing literature, we find that larger exporters react more at the price margin and less at the volume margin. Quantitatively, the heterogeneity is again substantial. Using coefficient estimates in column (1) we find that a firm exporting only one product is predicted to increase its volume by 5% following a 10% depreciation of its currency. By contrast, for a firm selling ten products the volume increases by only 2%.

6 Trade policy, market power and pricing-to-market

In this section, we focus on the predictions of our model relating trade policy instruments to the extent of PTM (proposition 3 for tariffs and corollary 1 for non-tariff measures).

6.1 Tariffs and pricing-to-market

Proposition 3 implies that tariffs by reducing the market share of foreign firms also reduce their incentive to engage in PTM. This is true as long at the strategic effect does not dominate, i.e. as long the firm market share is sufficiently small.[17] In order to test these predictions, we use data on applied tariffs that vary across product-destination cells but are time invariant and interact them with the exchange rate.[18] The corresponding estimating equation is as follows:

$$\ln \mathrm{UV}_{fdpt} = \beta_1 \ln e_{odt} + \beta_2 \left[\ln e_{odt} \times \ln(1 + t_{odp})\right] + \gamma \mathbf{x_{dt}} + \mathbf{FE}_{ot} + \mathbf{FE}_{fdp} + \varepsilon_{fdpt} \qquad (18)$$

where t_{odp} is the average tariff imposed by destination country d on product p imported from origin o over the period. In robustness checks we include additional interaction terms between $\ln e_{odt}$ and firm performance indicators, or alternatively between $\ln e_{odt}$ and destination, origin or product dummies. We expect β_2 to be negative, an increase in import tariffs acts as an aggregate decrease in productivity this is turn decreases the market share of all exporters of country o and dampens their PTM. Alternatively, we re-estimate equation (18) on two different samples based on the market share of the firm in the destination market. We expect β_2 to be more negative for firms with a low market share (i.e. when the strategic effect is likely to be negligible).

Table 6 reports the results. Column (1) report results using equation (18). In column (2) we further control for firm performance including an interaction term between the exchange rate variable and the number of products (considering other measures of performance leads to similar results as shown later). In both cases, we find strong support for proposition 3, i.e. tariffs decrease the elasticity of firm export prices to exchange rate.

[17] Note that within-firm and selection effects play in the same direction, as remaining firms do more PTM because they have more market power, while smaller firms, which do less PTM, exit, raising aggregate PTM by a composition effect.

[18] We do so for two reasons. First, we are not interested in the effects of tariffs variations but rather in the effect of high (average) tariff rates on the extent of PTM. Second, tariffs data contain many missing values and using average values allows maximizing the number of observations.

Are the effects of trade policy quantitative large? Faced with a 10% depreciation of its home currency, an exporter selling a product tariff-free in a given destination would raise his home-currency price by 1.7% (column 1). Faced with the same depreciation on a destination where he sold the same product with a 10% tariff he would raise it by only 1%. When the tariff reaches 20-25% there is no significant PTM left anymore.

Columns (3) and (4) report results for markets with a low (in the first quartile) versus high (in the last quartile) level of import tariffs, where the first quartile corresponds to zero tariffs. Results are consistent. In destinations-products characterized by zero tariff firms increase their price by 2% following a 10% depreciation. By contrast, in markets with high tariffs no significant pricing-to-market is detected. Taking PTM as an indication of firms' market power, these results suggest strong effects of trade policy on market structure.

In columns (5) and (6) we split the sample according to the firm's market share. Market shares are computed over total imports of destination d in product p, obtained from BACI (CEPII), where a product is defined at the 4-digit level.[19] As previously, we consider market shares from the beginning of the period. The coefficient on the interaction term between exchange rate and tariffs is negative and significant in both columns (5) and (6), but it is more than three times larger (in absolute value) in the sample of small firms. Thus, in line with the AB model's prediction, tariffs unambiguously reduce PTM only for small firms, which have a negligible impact on the destination's price index.

Table 6: Import Tariffs and pricing-to-market

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. var. ln unit values						
Variable			Tariffs		Market share	
Subsample			Low	High	Low	High
ln(RER)	0.171^a	0.026	0.192^a	-0.005	0.382^a	0.144^a
	(0.015)	(0.027)	(0.021)	(0.020)	(0.091)	(0.018)
ln(dest. GDP)	0.142^a	0.146^a	-0.001	0.251^a	0.103	0.213^a
	(0.024)	(0.024)	(0.046)	(0.035)	(0.108)	(0.032)
ln RER × ln(tariff+1)	-0.729^a	-0.718^a			-2.300^a	-0.644^a
	(0.111)	(0.112)			(0.651)	(0.133)
ln RER × ln(# product$_{t0}$)		0.060^a				
		(0.010)				
Observations	728839	728839	333059	186026	180884	191904
Adj. R^2	0.039	0.039	0.024	0.076	0.029	0.084

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses, except in column (6) in which standard errors are bootstrapped. All estimations include origin-year dummies and firm-product-destination fixed effects. High and Low means respectively above the third quartile or below the first quartile of the corresponding variable. Market share denotes the market share of the firm in the total imports of the destination market of a given HS4 product.

[19] http://www.cepii.fr/CEPII/en/bdd_modele/presentation.asp?id=1. Note that ideally, we would like to include the entire market sales, including the sales of destination d itself, but we do not observe these.

6.2 Non-tariff measures and pricing-to-market

Until now we have focused exclusively on import tariffs. Should we expect the effect of trade policy on PTM to be the same across policy instruments? How trade-policy instruments shape the adjustment of producer prices to exchange-rate changes depends on how they affect market power at the level of the firm. As argued in the introduction, tariffs and non-discriminatory NTMs importantly differ in this respect; their impact of PTM is thus expected to differ.

As stated in corollary 1, contrary to tariffs, non-discriminatory NTMs (e.g. sanitary and technical regulations) force the smallest firms out, raising the market share of all remaining ones (domestic and foreign alike) and raising their incentive to engage in PTM. Testing this prediction using data on ad-valorem of non-tariff measures is equivalent to estimating the following equation:

$$\ln \mathrm{UV}_{fdpt} = \beta_1 \ln e_{odt} + \beta_2 \left(\ln e_{odt} \times \mathrm{NTM}_{dp}\right) + \gamma \mathbf{x_{dt}} + \mathbf{FE}_{ot} + \mathbf{FE}_{fdp} + \varepsilon_{fdpt} \tag{19}$$

where NTM_{dp} the ad-valorem equivalent of non-tariff measures imposed by destination d on product p. Contrary to tariff, NTMs are recorded in the raw data as "MFN", i.e. applying to all origin countries.[20] We expect non-discriminatory non-tariff measures to have the opposite effect as tariffs, that is, they decrease competition for incumbents, raise market power, and should therefore increase PTM ($\beta_2 > 0$).

Tabel reports OLS estimates from equation(19). The results corroborate our hypothesis. PTM is significantly stronger quantitatively in markets with high levels of NTMs (column 1).

Results remain very similar when controlling for firm performance (column 2) or when bootstraping the standard errors to account for the fact that the NTMs variables are estimated (column 3). Importantly, results hold when splitting the sample according to the level of NTMs (columns 4 and 5). The effects are also quantitatively significant – albeit slightly smaller than in the case of tariffs. The coefficients in columns (1) to (3) suggest that moving from zero to a 10% ad valorem equivalent raises the price elasticity to exchange rate by half from 9% to 14%. Effects of similar magnitude are found when splitting the sample in columns (4) and (5).

6.3 Robustness Checks

Table 14 in appendix tests the robustness of the results for tariffs on PTM obtained in table 6. In columns (1) and (2) we use alternative measures of firms' performance. We show that the results are unaffected. In columns (3) to (5) we include additional interaction terms between the exchange rate and (i) destination group dummies; (ii) origin country dummies; and (iii) product dummies (as in Table 4). Again results are robust. This clearly confirms that we are identifying the effect of market-specific trade policy, rather than the role of other country or product specific determinants of pricing-to-market. Finally, in columns (6) and (7) we use export volume as an alternative dependent variable. Just like tariffs decrease the exchange rate elasticity in the case of price, they magnify the elasticity in the case of export volumes.

[20] Whereas applied tariffs are specific to origin-destination dyads, most non-tariff measures, in particular SPS and TBT regulations, are imposed on an "MFN" basis, i.e. specific to a destination and not a dyad. For instance, a maximum residual level of pesticides in horticulture products applies to *all* imports, not just to imports from a particular country, and, unlike a tariff, will not be relaxed in the presence of a preferential trade agreement.

Table 7: Non-tariff measures and pricing-to-market

	(1)	(2)	(3)	(4)	(5)
Dep. var. ln unit values					
Variable				NTMs	
Subsample				Low	High
ln(RER)	0.095^a	0.095^a	0.076^b	0.092^a	0.181^a
	(0.017)	(0.015)	(0.033)	(0.020)	(0.030)
ln(dest. GDP)	0.112^a	0.112^a	0.113^a	0.169^a	-0.056
	(0.035)	(0.042)	(0.035)	(0.042)	(0.061)
ln RER × NTM AVE	0.255^a	0.255^a	0.254^a		
	(0.041)	(0.008)	(0.041)		
ln RER × ln(# product_{t0})			0.008		
			(0.014)		
Observations	192620	192620	192620	141404	48910
Adj. R^2	0.054	0.054	0.054	0.045	0.096

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses, except in column (3) in which standard errors are bootstrapped. All estimations include origin-year dummies and firm-product-destination fixed effects. High and Low means respectively above the third quartile or below the first quartile of the corresponding variable.

Table 15 in appendix performs a series of robustness exercises for the effect of NTMs on PTM. In all columns, we include both trade policy measures simultaneously (this decreases importantly the number of observations compared to tariff regressions, which is why we take Table 7 as our baseline). We sequentially control for alternative indicators of firm size (columns 1 and 2), add interaction terms between the exchange rate and origin, destination or product dummies (columns 3 to 5). The results are extremely stable. Finally, in columns (6) and (7) we use export volumes as an alternative dependent variable. Export volumes react in a symmetric way compared to unit values, being more elastic to exchange-rate changes when tariffs are high and less elastic when NTMs are high. However, the coefficients on our interaction terms are significant only in the case of tariffs.

6.4 Extension – The role of market shares

The effect of trade policy on pricing-to-market is supposed to go through changes in market shares at the firm-level. In this section, we explore this idea further and try to directly estimate the effect of trade policy instruments on PTM *through variations in the firms' market share.*

Table 8 column (1) reports the results of a preliminary analysis where we include in our baseline specification (equation (15)) an interaction term between the firm average market share in destination d, product p and the real exchange rate. Again, we define a sector as a 4-digit product and we compute market share as the share of firms' export in the total imports of the destinations for that HS4 product. The coefficient on the interaction term between market shares and the real exchange rate is positive and statistically significant at the 5 percent level, a result consistent with Amiti *et al.* (2014).

Average market shares are however potentially endogenous (influenced by prices). We instrument the interaction between market shares and exchange rate by interaction terms between the exchange rate and our two trade policy measures. We expect the firms with higher market shares due to lower tariffs or higher non-tariff barriers to adjust more their prices when exchange rate varies. Table 16 in the appendix shows that this is indeed the case. Firm-destination-product (HS4) specific market shares are positively correlated with NTMs and negatively with applied tariffs.[21]

Column (2) of Table 8 uses the interaction terms between trade policy instruments and the real exchange rate as instrumental variables. We find a positive and significant effect of market share on firm-market specific degree of pricing-to-market, when instrumented by trade policy.

Column (3) uses product-destination rather than firm-product-destination specific market shares. The coefficients are more precisely estimated in this case, which is to be expected as trade policy instruments affect all exporters from a given origin country in a symmetric way. Note that in the first stage, only the interaction with tariffs is statistically significant. This might be due to the fact that tariffs vary more, as they are *de facto* bilateral, contrary to NTMs.

Note that this exercise has a number of drawbacks. First, we cannot directly instrument the firm market share with trade policy instruments due to the inclusion of firm-destination-product fixed effects–only the interaction can be instrumented. Second, it is not clear how exactly market power should be measured and in particular what is the relevant "market" on which to compute market shares. Third, it is possible that our firm-level custom data and UN-COMTRADE data do not perfectly match, making the computation of market shares problematic. For all these reasons, we consider this exercise as a complement to our baseline results shown in Table 6.

7 Concluding remarks

Our objective in this paper was double. First, we set out to explore in some depth the determinants of pricing-to-market for a sample of developing-country exporters. Second, we proposed to put the PTM literature "on its head", starting from the assumption that PTM reflects market power and, based on that assumption, exploring how different trade-policy instruments affect market power using exporter adjustment to exchange-rate fluctuations as the identification mechanism. The size and dimensionality of our multi-country, firm-level dataset allowed us to filter out many confounding influences with a powerful array of fixed effects.

As to the first objective, we were able to confirm results obtained so far largely on industrial-country data in a more general setting and with a powerful identification. We found that developing-country exporters in our sample typically absorb about ten to fifteen percent of the effect of currency fluctuations, passing through the remaining 85-90%. There was surprisingly little variation in this split, even though our sample spanned several continents and included countries at different levels of development and integration in global value chains. Also in accordance with the literature, we found that, on the basis of various proxies for firm size (and hence performance), PTM clearly rises with exporter size. Quantitatively, the effect is sizable; for instance, in reaction to a given home-currency depreciation, a firm that exports ten products (overall) would raise its home-currency price three times more than one that exports a single

[21] Trade policy variables also have a direct impact on unit values, but this effect is captured by the firm-product-destination fixed effects, given that we use time-invariant policy measures.

Table 8: Market share, trade policy and pricing-to-market

Dep. var: ln unit value	(1)	(2)	(3)
ln(RER)	0.101^a	-0.296	-0.189
	(0.018)	(0.203)	(0.129)
ln(dest. GDP)	0.160^a	0.218^a	0.189^a
	(0.037)	(0.074)	(0.072)
ln RER × market share$_{fdp}$	0.740^b	29.976^b	
	(0.354)	(14.883)	
ln RER × market share$_{odp}$			1.595^a
			(0.591)
First stage (dep. var.: ln RER × market share)			
ln RER × ln(tariff+1)		-0.025^a	-0.491^a
		(0.007)	(0.139)
ln RER × NTM AVE		0.001	0.110
		(0.003)	(0.101)
ln RER		0.015^a	0.198^a
		(0.001)	(0.028)
ln(dest. GDP)		-0.002^a	-0.012^c
		(0.001)	(0.007)
Observations	121607	121607	120036
Hansen overid stat.	-	2.4	0.40
- p-value	-	0.12	0.53
F-stat excl. instruments	-	7.0	7.5

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. All estimations include origin-year dummies and firm-product-destination fixed effects.

product.

As to the second objective, we found that tariffs in a destination market reduce the extent of PTM by exporters selling on that destination market. This is in accordance with the rent-shifting effect of tariffs in traditional oligopoly theory. Moreover, we found that the differential impact of higher tariffs on heterogeneous firms was consistent with the logic of a model where variable markups derived from a nested two-level CES structure. We found that non-tariff measures have the opposite effect, being associated with more PTM. This is again consistent with theory if non-tariff measures raise costs for all firms alike (domestic and foreign), inducing the exit of the smaller ones and consequently larger market shares (and market power) for those that stay.

While the argument that trade policy affects competition is an old one, there is to this day little firm-level evidence of how it does so and, in particular, how the effects of different trade-policy instruments play out. Our results suggest that useful information can be generated in this regard from the analysis of exporter adjustment to exchange-rate fluctuations.

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Appendix 1: theory

Derivation of Γ

Consider a nested two-level CES demand system where foreign consumers have CES preferences over a continuum of products indexed by p, and within products over a discrete set of varieties, each produced by a firm, indexed by f. The elasticity of substitution is η between products and ρ between varieties, with $1 < \eta < \rho$. In order to keep the notation in this appendix consistent with that in the body of the paper but nevertheless reasonably light, we index varieties by fp (firm-product) while omitting dt (destination-time) and stars, all magnitudes here being measured in the destination's currency. Expressions derived in this appendix all refer to a destination-year cell. The upper-level aggregate, consumption, is

$$Q = \left[\int_p Q_p^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}} \tag{20}$$

with price index

$$P = \left[\int_p P_p^{1-\eta} \right]^{\frac{1}{1-\eta}}. \tag{21}$$

The CES aggregator of varieties into products is

$$Q_p = \left[\sum_f q_{fp}^{\frac{\rho-1}{\rho}} \right]^{\frac{\rho}{\rho-1}} \tag{22}$$

with corresponding product-level price index

$$P_p = \left[\sum_f p_{fp}^{1-\rho} \right]^{\frac{1}{1-\rho}}. \tag{23}$$

Given those, the inverse demand for product p is

$$\frac{P_p}{P} = \left(\frac{Q_p}{Q} \right)^{\frac{-1}{\eta}} \tag{24}$$

and the optimal pricing rule is, as usual,

$$p_{fp} = \mu_{fp}\, c_{fp} = \left(\frac{\varepsilon_{fp}}{\varepsilon_{fp} - 1} \right) c_{fp} \tag{25}$$

where, after some algebraic manipulation assuming quantity competition,

$$\varepsilon_{fp} = \frac{\rho\eta}{\rho s_{fp} + \eta(1 - s_{fp})} \tag{26}$$

with s_{fp} defined in (3) in the text. Using these,

$$\begin{aligned}\Gamma_{fp} &= -\frac{d\ln\mu_{fp}}{d\ln\varepsilon_{fp}}\frac{d\ln\varepsilon_{fp}}{d\ln s_{fp}}\bigg|_{P_p^\star\text{constant}} = \left(\frac{1}{1-\varepsilon_{fp}}\right)\frac{\rho-\eta}{\rho\eta}\,s_{fp} \\ &= \left\{\frac{s_{fp}}{1-[(1-s_{fp})/\rho]-s_{fp}/\eta}\right\}\left(\frac{1}{\rho}-\frac{1}{\eta}\right). \end{aligned} \tag{27}$$

Derivation of (6)

Under CES preferences, the demand facing variety p from firm f is given by

$$\frac{q_{fp}}{Q_p} = \left(\frac{p_{fp}^\star}{P_p^\star}\right)^{-\rho} \tag{28}$$

so

$$s_{fdpt}^\star = \frac{p_{fp}^\star}{P_p^\star}\frac{q_{fp}}{Q_p} = \left(\frac{p_{fp}^\star}{P_p^\star}\right)^{1-\rho}. \tag{29}$$

Log-differentiating (29) gives the result directly.

Appendix 2: additional tables

Table 9: Summary statistics: Trade protection

Country	NTM		MFN Tariff		Country	NTM		MFN Tariff	
	mean	s.d.	mean	s.d.		mean	s.d.	mean	s.d.
Argentina	0.16	0.43	0.12	0.07	Lebanon	0.01	0.09	0.07	0.09
Austria	0.11	0.35	0.04	0.04	Lithuania	0.17	0.38	0.04	0.04
Bangladesh	0.17	0.45	0.17	0.11	Luxembourg	0.19	0.46	0.04	0.04
Bolivia	0.01	0.11	0.09	0.03	Madagascar	0.04	0.24	0.11	0.06
Brazil	0.18	0.41	0.13	0.07	Mauritius	0.03	0.29	0.05	0.09
Bulgaria	0.13	0.36	0.04	0.04	Mexico	0.09	0.36	0.15	0.14
Burundi	0.13	0.44	0.15	0.10	Morocco	0.04	0.25	0.24	0.24
Cambodia	0.03	0.24	0.14	0.11	Namibia	0.08	0.41	0.08	0.11
Chile	0.02	0.27	0.07	0.02	Paraguay	0.02	0.19	0.11	0.07
China	0.26	0.53	0.13	0.09	Peru	0.01	0.11	0.09	0.05
Colombia	0.02	0.15	0.12	0.07	Poland	0.13	0.33	0.04	0.04
Czech Republic	0.15	0.35	0.04	0.04	Senegal	0.01	0.12	0.11	0.07
Ecuador	0.05	0.26	0.11	0.08	Slovak Republic	0.16	0.37	0.04	0.04
Egypt, Arab Rep.	0.17	0.56	0.19	1.15	Slovenia	0.17	0.38	0.04	0.04
Finland	0.18	0.37	0.04	0.04	South Africa	0.05	0.23	0.08	0.11
Hungary	0.03	0.33	0.04	0.04	Sri Lanka	0.19	0.50	0.10	0.12
India	0.19	0.52	0.22	0.13	Syrian Arab Republic	-0.01	0.27	0.15	0.18
Indonesia	0.07	0.33	0.08	0.10	Tanzania	0.00	0.10	0.14	0.10
Japan	0.11	0.35	0.03	0.05	Tunisia	0.05	0.24	0.26	0.24
Kazakhstan	0.14	0.39	0.05	0.04	Uganda	0.21	0.48	0.10	0.09
Kenya	0.18	0.47	0.14	0.11	Uruguay	0.00	0.09	0.11	0.07
Lao PDR	0.19	0.45	0.09	0.08	Venezuela, RB	0.07	0.30	0.12	0.07
Latvia	0.17	0.37	0.04	0.04	**Average**	0.10	0.33	0.11	0.11

Source: see section 2.3.

Table 10: Exchange rates and unit values, robustness

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dep. var.: ln(unit value)							
Exporting countries	All	Excl. Mexico	EAC	MENA	Bangladesh	Chile	Mexico
ln(RER)	0.109^a	0.108^a	0.070	0.070^b	0.058^c	0.137^a	0.111^a
	(0.014)	(0.014)	(0.112)	(0.035)	(0.030)	(0.015)	(0.032)
ln(GDP dest.)	0.113^a	0.069^a	0.143	0.084	0.086	0.044	0.312^a
	(0.027)	(0.025)	(0.156)	(0.058)	(0.060)	(0.028)	(0.093)
Observations	1123384	562255	54684	182035	127600	197936	561129

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. All estimations include *origin-product-year* and firm-product-destination fixed effects.

Table 11: Exchange rates and export volumes

Dep. var.: ln(export volume)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Exporting countries	All	Excl. Mexico	EAC	MENA	Bangladesh	Chile	Mexico
ln(RER)	0.143^{a}	0.182^{a}	-0.018	0.309^{a}	0.183^{c}	0.179^{a}	0.056
	(0.026)	(0.034)	(0.129)	(0.076)	(0.098)	(0.044)	(0.040)
ln(GDP dest.)	1.394^{a}	1.218^{a}	1.479^{a}	1.336^{a}	1.143^{a}	1.086^{a}	2.270^{a}
	(0.052)	(0.058)	(0.200)	(0.098)	(0.179)	(0.080)	(0.105)
Observations	1123384	562255	54684	182035	127600	197936	561129
Adj. R^2	0.076	0.027	0.017	0.045	0.028	0.007	0.142

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. All estimations include origin-year dummies and firm-product-destination fixed effects.

Table 12: Exchange rates and export volumes, robustness

Dep. var.: ln(export volume)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Exporting countries	All	Excl. Mexico	EAC	MENA	Bangladesh	Chile	Mexico
ln(RER)	0.144^{a}	0.212^{a}	-0.061	0.271^{b}	0.262^{c}	0.180^{a}	0.021
	(0.040)	(0.049)	(0.196)	(0.111)	(0.141)	(0.059)	(0.062)
ln(GDP dest.)	1.330^{a}	1.153^{a}	0.185	0.856^{a}	1.861^{a}	1.222^{a}	2.046^{a}
	(0.083)	(0.086)	(0.308)	(0.151)	(0.276)	(0.111)	(0.186)
Observations	1123384	562255	54684	182035	127600	197936	561129

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. All estimations include *origin-product-year* and firm-product-destination fixed effects.

Table 13: Firm heterogeneity and export volumes

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. var.	ln unit value					
ln(RER)	0.508^a	0.820^a	-0.186^a			
	(0.045)	(0.058)	(0.071)			
ln(dest. GDP)	1.368^a	1.373^a	0.056^a			
	(0.051)	(0.052)	(0.020)			
ln(RER) × ln(# prod_{t0})	-0.136^a			-0.251^a		
	(0.012)			(0.028)		
ln(RER) × ln(# dest_{t0})		-0.248^a			-0.339^a	
		(0.018)			(0.041)	
ln(RER) × ln(export val_{t0})			0.019^a			-0.136^a
			(0.004)			(0.019)
Observations	1123290	1123290	1123289	1132939	1132939	1132938
Firm-product-destination FE	Yes	Yes	Yes	Yes	Yes	Yes
Origin-year FE	Yes	Yes	Yes	No	No	No
Origin-destination-product-year FE	No	No	No	Yes	Yes	Yes

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. ln(# prod_{t0}), ln(# dest_{t0}) and ln(export val_{t0}) are the number of products, the number of destinations and the total export value of the firm during the first year it appears in the sample.

Table 14: Tariffs and Pricing-to-market: robustness

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dep. var:	ln unit value					ln export volume	
ln(RER)	-0.024	-0.340[a]	0.091[a]			0.211[a]	0.751[a]
	(0.037)	(0.096)	(0.033)			(0.038)	(0.065)
ln(GDP dest.)	0.148[a]	0.141[a]	0.137[a]	0.144[a]	0.137[a]	1.142[a]	1.126[a]
	(0.024)	(0.024)	(0.025)	(0.025)	(0.025)	(0.061)	(0.060)
ln(RER)×ln(1 + tariff)	-0.757[a]	-0.724[a]	-0.573[a]	-0.574[a]	-0.591[a]	0.169	0.127
	(0.112)	(0.112)	(0.111)	(0.116)	(0.119)	(0.303)	(0.306)
ln rer × ln(# dest_{t0})	0.074[a]						
	(0.013)						
ln(RER) × ln(export val_{t0})		0.032[a]					
		(0.006)					
ln(RER) × ln(# product_{t0})			0.071[a]	0.057[a]	0.056[a]		-0.224[a]
			(0.010)	(0.010)	(0.012)		(0.020)
Additional interactions							
ln rer × destination group	No	No	Yes	Yes	Yes	No	No
ln rer × origin	No	No	No	Yes	Yes	No	No
ln rer × HS2	No	No	No	No	Yes	No	No
Observations	728839	728839	728839	728839	728839	728839	728839

[c] significant at 10%; [b] significant at 5%; [a] significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. All estimations include origin-year dummies and firm-product-destination fixed effects.

Table 15: NTMs and Pricing-to-market: robustness

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dep. var:	ln unit value					ln export volume	
ln(RER)	-0.211[a]	-0.078	0.130[a]			0.150[a]	0.401[a]
	(0.049)	(0.112)	(0.048)			(0.051)	(0.092)
ln(GDP dest.)	0.155[a]	0.152[a]	0.165[a]	0.164[a]	0.162[a]	1.334[a]	1.324[a]
	(0.037)	(0.037)	(0.037)	(0.038)	(0.038)	(0.106)	(0.106)
ln(RER)×ln(1 + tariff)	-0.577[a]	-0.529[a]	-0.410[b]	-0.438[b]	-0.152	0.875[b]	0.815[b]
	(0.156)	(0.156)	(0.165)	(0.176)	(0.179)	(0.409)	(0.410)
ln(RER)×NTM AVE	0.294[a]	0.262[a]	0.243[a]	0.234[a]	0.161[a]	-0.165	-0.159
	(0.042)	(0.041)	(0.041)	(0.041)	(0.052)	(0.119)	(0.120)
ln rer × ln(# dest_{t0})	0.123[a]						
	(0.016)						
ln(RER) × ln(export val_{t0})		0.012[c]					
		(0.007)					
ln(RER) × ln(# product_{t0})			0.005	0.010	0.005		-0.105[a]
			(0.013)	(0.013)	(0.015)		(0.029)
Additional interactions							
ln rer × destination group	No	No	Yes	Yes	Yes	No	No
ln rer × origin	No	No	No	Yes	Yes	No	No
ln rer × HS2	No	No	No	No	Yes	No	No
Observations	183473	183473	183473	183473	183473	183473	183473

[c] significant at 10%; [b] significant at 5%; [a] significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. All estimations include origin-year dummies and firm-product-destination fixed effects.

Table 16: Market share, trade policy and pricing-to-market

Dep. var: market share$_{fdpt}$	(1)	(2)	(3)
ln(1 + tariff)	-0.012^{a}	-0.003^{b}	-0.004^{b}
	(0.001)	(0.001)	(0.001)
NTM AVE	0.001^{c}	0.001^{b}	0.001^{b}
	(0.001)	(0.001)	(0.001)
Observations	95356	95356	95356
Adj. R^2	0.051	0.305	0.314
Destination FE	Yes	Yes	No
Product FE	No	Yes	Yes
Origin-destination FE	No	No	Yes

^{c} significant at 10%; ^{b} significant at 5%; ^{a} significant at 1%. Robust standard errors clustered by product-origin-destination are in parentheses.