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Trade Policy and Market Power: Firm-level Evidence

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Abstract

This paper identifies the effect of trade policy on market power through new data and a new identication strategy. 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, merged with destination-product specific information on tariffs and non-tariff barriers. We identify market power by observing how exporting firms price discriminate across markets in reaction to variations in bilateral exchange rates. Pricing-to-market is prevalent in all regions of our sample, even among small firms, although it is increasing in firm size, in accordance with theory. More importantly, we find that the effect of non-tariff measures is not isomorphic to that of tariffs: the pricing-to-market behavior we observe suggests that, while tariffs reduce the market power of foreign firms through classic rent-shifting effects, non-tariff measures alter market structure and reinforce the market power of non-exiting firms, domestic and foreign ones alike.

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1 Introduction

While the argument that trade policy affects competition is an old one, there is to date little systematic evidence on how trade policy affects market power at the level of the firm and, in particular, how the effects of different trade-policy instruments play out. Indeed, in a heterogeneous-firms setting, non-tariff measures (NTMs)—a widely used class of trade-policy instrument including e.g. technical or sanitary regulations—affect market structure in ways that are not isomorphic to tariffs: Rather than merely shifting rents from foreign to domestic firms, they also change market structure by affecting firms differently depending on their size. For instance, a regulation may inadvertently create barriers to entry that generate a dominant position. Evidence of such effects of tariffs and NTMs is scarce, however, because assessing the market power of firms and how it relates to trade policy requires estimating indicators such as price-cost margins or market shares in well-identified markets. Both the identification of these measures and the definition of a market (and its structure) are problematic, which is why research has so far limited itself to specific sectors, countries and trade policy instruments.

This paper follows a different route and identifies market power and how it is affected by trade policy using insights from the pricing-to-market (PTM) literature. We rely on a recent and growing literature that establishes, theoretically and empirically, that the extent of firm-level PTM relates to firm size, efficiency and, ultimately, market power.[1] PTM implies that a firm faced with a cost shock will adjust its price differentially across destination markets, depending on the price elasticity of demand it perceives on each of them, which in turn depends on its market power. However, neither cost shocks nor market shares are observed directly at the firm level.[2] Our identification strategy goes around this by relating firm-level export prices (on which we have data) to bilateral exchange rate shocks.

Consider a firm faced with a series of bilateral exchange rate shocks, one on each of its destination markets. If it passed through the entirety of each shock onto consumer prices, its producer price in the home currency would remain the same irrespective of destination; there would be no PTM. However, with incomplete pass-through, a fraction of each shock is absorbed by the firm into its producer price; as a result, the uniqueness of the producer price breaks down, giving rise instead to a series of differentiated ones by destination; in other words, the firm prices to market. Like in the case of a symmetric cost shock, the reaction of a firm's producer price to an exchange rate shock tells us something about the elasticity of demand it perceives on its destination market—i.e., about its unobserved market power. Specifically, a firm with strong market power in a given destination (hence a low perceived price elasticity of demand) will absorb a large fraction of the bilateral exchange rate shock, and conversely. Because destination markets vary in terms of their trade policy, we can then infer the effect of trade policy on market structure by observing the pricing behaviour of firms exporting there, without directly needing to observe their market shares or any other measure of market power. As our identification relies on an interaction term (between bilateral exchange rate shocks and destination-product-specific trade-policy measures) we can control for a large set of confounding influences at the level of both

[1] See for example Krugman (1994), Feenstra, Gagnon and Knetter (1996), Atkeson and Burstein (2008), Berman, Martin and Mayer (2012), Auer and Schoenle (2016), or Amiti, Istkhoki and Konings (2014).

[2] Market shares are not even observed at the more aggregated product level, as domestic firm sales are typically not observed.

the firm and the market with a powerful array of fixed effects.

We start by presenting a simple theoretical framework based on Atkeson and Burstein (2008) which we use to derive testable predictions on the extent of PTM by exporting firms and how it relates to their market power and the trade policy environment they face in destination markets. We consider both non-tariff measures and tariffs. Specifically, we show that NTMs applied in non-discriminatory fashion—that is, in compliance with the WTO's "national treatment" clause, whereby imported and domestically-produced products must be treated alike—either have no effect on PTM or raise it for incumbents if they induce the exit of smaller firms, e.g. through higher fixed costs.[3] As for tariffs, the model suggests that the presence of import tariffs in some destination markets has an ambiguous effect on the extent of firm-level PTM.

We then perform an empirical test for the model's predictions. We rely on a large multi-country firm-level dataset obtained from customs administrations in twelve developing countries, ranging from low income in the case of Uganda to OECD in the case of Mexico. We combine the firm-level data with destination-product specific data on bilateral applied tariffs as well as non-tariff measures. The latter covers a wide range of measures ranging from sanitary and phytosanitary measures to technical barriers to trade. Pooling together firm-level data from several developing countries is a first and lends itself to a more robust and systematic exploration of the effect of trade policy on competition and market power. The high dimensionality of the data enables us to go beyond existing work in accounting for unobserved heterogeneity as well as for the possible endogeneity of tariffs and non-tariff measures to the market structure within sector and destinations. In particular, we are able to control for any firm characteristics which might affect their reaction to exchange rate variations and may be correlated with trade policy measures. The fact that our unit of observation is a firm rather than a product or a country also implies that our estimates are unlikely to be driven by large firms, which is important given the skewness of the firm-size distribution (Mayer and Ottaviano, 2007; Freund and Pierola, 2012).

We make two contributions. First, we show that PTM is prevalent in our entire sample, implying that even small firms in developing countries do it; Nevertheless, large firms do more of it than smaller ones in our data as in other recent studies. Our estimate of the elasticity of home-currency FOB export prices to exchange rates implies pass-through at a rate quite similar to what existing studies have found for industrial countries. Faced with a 10% bilateral exchange rate depreciation on a given market (a reduction in the value of the foreign currency), firms in our sample cut their home-currency price for export to that market by around 1.5% on average, implying incomplete (85%) although arguably still high exchange rate pass-through to export prices. Interestingly, this rate remains quite stable across exporting countries and regions.

Second, our approach allows us to highlight the effect of trade policy on market structure in a non-conventional way. In accordance with the model prediction, when exporters face non-tariff measures on their destination markets, they do more PTM. If market i is affected by an NTM with a 10% ad-valorem equivalent (AVE), the wedge between the producer price for export to i and that for export to j following again a 10% exchange-rate shock will be higher by a third, from 1% to 1.3%. Tariffs have the opposite effect. Instead of doing more PTM, exporters faced with

[3] 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.

tariffs on their destination markets do significantly *less*, as if tariffs were robbing them of some of their market power through rent-shifting effects; a 10% tariff thus reduces the elasticity of the home-currency export price to the exchange rate by a third, from two percent to 1.3 percent. To put things differently, consider an initial situation where a firm prices symmetrically on all markets. In the absence of tariffs and NTMs, a 10% bilateral exchange rate shock on market i will create, ceteris paribus, a 2% wedge between the producer price of a given good shipped by that firm to i and the same good shipped by the same firm to market j. In the presence of a 10% tariff on market i, the same shock will create only a 1.3% wedge. When tariffs reach 30-35% no significant adjustment is detected anymore in the home-currency export price, implying full pass-through. Considered jointly, the differential effects of tariffs and NTMs on PTM can be large. In markets with zero tariffs but with NTMs with large AVEs, we find that exchange rate pass-through drops dramatically from around 90% to 60%.

What could account for this differential effect of tariffs and NTMs? In the absence of timewise variation in NTMs, we cannot test directly their effects on prices or quantities, but only indirectly through the interaction with exchange-rate shocks. A plausible conjecture is that NTMs have rent-shifting effects eroding the market power of foreign exporters in favor of domestic ones only if they are discriminatory like tariffs. If they are not, they generate compliance costs for *all* producers that induce the exit of the smallest ones, raising the market share of all remaining ones, including foreign exporters. In that case, the market power of foreign exporters is enhanced by NTMs. The notion that NTMs may affect market structure in terms of large vs. small firms more than home vs. foreign is quite new, as most of the literature on NTMs has assumed so far that they were mere surrogates for tariffs (see e.g. Feenstra 1984, Deardorff and Stern 1997, Baldwin 1989, Leamer 1990, Anderson and Neary 1994, Kee, Nicita and Olarreaga 2009, Carrere and de Melo 2011, or Cadot and Gourdon 2014a). Thus, our approach provides a test of whether NTMs are, on average, applied in discriminatory fashion, an important and largely unexplored policy issue. Our results are robust to a battery of sensitivity checks. In particular, they still hold when controlling for all changes in firm-level costs which might be correlated with exchange rate movements through the inclusion of firm×product×year fixed effects.

Finally, although we do not have direct measures of market share (which is why we use exchange rate pass-through as our identification mechanism), as a robustness test, we follow Amiti *et al.* (2014) and proxy the market share of firm f on destination market d for product p as the share of firm f in destination d's total imports of product p. Using tariffs and NTMs as instruments for firm-specific market shares in a given product-destination cell confirms our claim that changes in the pricing behavior of exporters reflects the effect of destination trade policy on market structure. We find that large market shares amplify PTM and that trade policy has a significant effect on market shares.

Our results have important policy implications. Indeed, with tariffs on many manufactured products down to low levels, non-tariff measures have taken on increasing importance in regional trade negotiations, whether in "mega-regionals" (e.g. TPP or TTIP) or existing ones such as ASEAN, which have tried to prevent them from replacing tariffs as trade barriers. Current policy approaches, which draw on the past ten years of literature on NTMs, consist essentially in quantifying their trade effects through ad-valorem equivalents, feeding them into computable general equilibrium models, and hoping to cut them down through inter-governmental negotiations. However, little progress has been achieved so far. While political-economy factors may well have contributed to the lack of progress, this paper argues that trade-based approaches may also be going down the wrong alley. Our results suggest that NTMs should be viewed as a competition-policy issue as much as a trade one. Considering them through this perspective would help shift the policy perspective away from doomed trade negotiations to more constructive approaches emphasizing cooperation between regulatory and antitrust agencies (within and between countries) in the design of regulations.

Our work relates to various strands of the literature. Our paper is close to a number of papers that examine the role of trade policy and trade reforms on market power and pricing behaviour. For instance, looking at the U.S. steel industry and using plant-level census data, Blonigen *et al.* (2013) find that, as suggested by traditional theory, quota-based protection increases the market power of domestic producers while tariff-based protection does not. Kim (2000) provide similar results based on Korean sector-level data. Antidumping protection, which takes the form of tariffs (or, sometimes, price agreements), has also received some attention, although results are ambiguous.[4] Papers by De Loecker *et al.* (forthcoming) and De Loecker, Fuss and Van Biesebroeck (2014) also provide evidence that the markups of domestic firms are affected by tariff reductions. These studies generally use measures of price-cost margins (PCM) at the firm level to identify market-power effects.

The main problem faced by this literature is that estimating markups over marginal costs requires the use of detailed data on costs and/or strong identifying assumptions.[5] Using market shares as a measure of market power is also problematic, both because one has to define what the appropriate "market" is and because, when studying exporters, the sales of domestic firms are typically not observed. Most of the papers cited above use very detailed firm-level data, allowing for a precise identification of the channels trough which trade policy affect markups; but this comes at a cost: these studies are typically limited to a single country—often a single industry—and focus exclusively on domestic producers. Using exchange rate variations to identify price discrimination and therefore market power across destinations has the advantage of requiring data on neither firm costs nor market shares. The idea that useful information on the extent of market power can be generated from price adjustments to exchange rate fluctuations is not new, as it was already present in early work such as Aw (1993), Goldberg and Knetter (1999) or Bernhofen and Xu (1999).[6] Our study, which is close in spirit to those papers, links the literature on PTM with that on trade policy and competition.

Finally, our paper also relates to a vast literature on the determinants of PTM. On the theory side, our approach follows Atkeson and Burstein (2008) and subsequent papers by Amiti, Itskhoki and Konings (2014) or Auer and Schoenle (2016). 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

[4]Blonigen *et al.* (2013) estimate a positive but mostly insignificant effect on market power. Konings and Vandebussche (2005) and Pierce (2011) find significant evidence of enhanced market power for EU and US firms respectively, while Nieberding (1999) and Rovegno (2013) find mixed results.

[5]See the recent survey by De Loecker and Van Biesebroeck (2015) which discusses the different methods (production-based or demand-based) which have been used to estimate mark-ups.

[6]See the survey by Goldberg and Knetter (1997).

domestic sales) implying different degrees of markup adjustment in response to firm-level cost shocks. Heterogenous mark-up adjustment across firms with different market shares can also be obtained in alternative models, featuring distribution costs (e.g. Corsetti and Dedola, 2005) or non-CES preferences such as quadratic preferences *à la* Melitz and Ottaviano (2008).[7] All these models generate PTM, the extent of which depends on the firms' market share, itself a correlate of productivity and size. Some yield an unambiguous relationship between size and pricing behaviour, where large firms perceive a lower elasticity of demand, which makes their markups more responsive to exchange rate movements.[8] 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.

The relationship between PTM and market power highlighted in the theory also appears as an empirical regularity. The empirical literature on the link between pass-through and firm size goes back at least to the work of Feenstra, Gagnon and Knetter (1996).[9] 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), Li, Ma and Xu (2015) and Chen and Juvenal (2016) provide similar evidence using respectively Brazilian, Chinese and Argentinean firm-level data. We confirm these results hold on a larger and more diversified set of countries – some low, some middle-income – and additionally show that trade policy deeply affects the reaction of export prices to exchange rate movements.

This remainder of the paper is structured as follows. The next section presents a summary model drawing on Atkeson-Burstein (2008) and derives three main predictions to guide the empirical analysis. Section 3 presents our multi-country firm-level dataset and the following sections test our predictions: Section 4 estimates the extent of PTM at the firm-level across the different countries of our sample, and how it varies across firms, and Section 5 examines the effect of trade policy. Finally, Section 6 concludes.

2 Theory

In this section we use a simple theoretical framework based on a variable-markup model à la Atkeson and Burstein (2008, henceforth AB) and derive several testable predictions on the extent of PTM and how it relates to market power and destination-market trade policy.

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 identify PTM through shocks on bilateral exchange rates rather than on firm-wide production costs. Exchange-rate shocks are equivalent to destination-specific shocks on the exporter's marginal cost expressed in destination-currency. The magnitude of the reaction of producers'

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

[8]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.

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

home-currency prices to these shocks determines the degree of pass-through: No reaction translates into complete pass-thought with *no* changes of home-currency prices, and accordingly, incomplete pass-through translates into *some* changes. In order to stay close to the empirics, 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 bilateral exchange rate shocks, which are asymmetric between destinations, implies PTM as it drives wedges between home-currency export prices to different destinations.

Suppose there are only two countries, home (origin) and foreign (destination). Let stars denote variables expressed in destination currency and hats denote log-changes. Let μ_{fdpt} be home firm f's markup over marginal cost when it sells product p to destination d at time t, c_{fpt} and $c^{\star}_{fdpt} = c_{fpt}/e_{odt}$ its marginal cost expressed in home and destination 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. Thus, when the home currency depreciates, e goes up.[10] Market structure is determined by a symmetric fixed cost F which will play a role in the analysis only when we consider the effect of trade-policy instruments. Home firm f's foreign-currency price is:

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

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 of firm f's home currency ($\hat{e}_{odt} > 0$) is equivalent to a negative shock on its marginal cost expressed in the destination currency ($\hat{c}^{\star}_{fpt} < 0$). 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. Like AB, we assume a two-level CES demand system with elasticities of substitution η between products and $\rho > \eta > 1$ between varieties. Let Γ_{fdpt} be the elasticity of home firm f's markup to its market share in destination d in a quantity-setting game:

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

The derivation for expression (4) is provided in the appendix. Note that Γ_{fdpt} is increasing in $s^{\star}_{fdpt}$, that $\Gamma(0) = 0$, and that $\Gamma(1)$ is 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 as a function of the log-change

[10] 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, p) couple is what it calls a 'variety'.

in its price relative to the log-change in destination d's sectoral price index:

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

Combining (5) and (6) and rearranging, the change in home firm f's destination-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_{fdpt}. Now, since $p_{fdpt} = e_{odt}\ p^{\star}_{fdpt}$, the degree of pass through measured as the change in firm f's home-currency price can be written, after some further rearrangement, as:

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

There are only two countries, home (origin) and foreign (destination), and consider the effect of a depreciation of the home currency, $\hat{e}_{odt} > 0$. Then $\hat{P}^{\star}_{dpt} < 0$ (destination d's sectoral price index, in foreign currency, goes down because imported varieties are now cheaper); 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 we have $\mid \hat{P}^{\star}_{dpt} \mid < \hat{e}_{odt}$, which drives two results. First, the term in parentheses in (9) is positive and as a result the degree of pass-through will depend positively on parameter λ_{fdpt}. Second, by (4), $0 < \lambda_{fdpt} < 1$ whenever $s^{\star}_{fdpt} > 0$; so (9) implies that $0 < \hat{p}_{fdpt} < \hat{e}_{odt}$; there is some pass-through, but it is not complete (pass-through would be complete with $\hat{p}_{fdpt} = 0$, i.e. 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 the pass-through parameter 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 exchange rate pass-through coefficient. By (5) and (9), given that $\Gamma(0) = 0$:

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

so pass-through is complete for very small firms. Similarly, for a very large firm ($s^{\star}_{fdpt} \to 1$), $\hat{p}^{\star}_{fdpt} \to \hat{P}^{\star}_{dpt}$ so, by (6), $\hat{s}^{\star}_{fdpt} \to 0$; as $\Gamma(1)$ is finite, again by (5) and (9) we have:

$$\lim_{s^{\star}_{fdpt}\to 1} \hat{p}^{\star}_{fdpt} = -\hat{e}_{odt} \quad \Rightarrow \quad \hat{p}_{fdpt} = 0. \tag{11}$$

Thus, both very small and very large firms keep their home-currency prices constant following a depreciation, i.e. tend to full pass-through, which is a non-monotone function of firm size/market share, as discussed in detail by Auer and Schoenle (2016).

Can we say anything more to guide the firm-level empirics? Consider a home firm f exporting a given variety of product p to two identical destinations, 1 and 2, with an iceberg trade cost $\tau > 1$ applying only to destination 2. Suppose now that firm f faces an identical exchange-rate shock on the two destination markets $\hat{e}_{ot} = \hat{e}_{odt}, d = 1, 2$ (its home currency depreciates by the same amount relative to the currencies of the two destination markets), inducing it to adjust its foreign-currency prices by $\hat{p}^{\star}_{f1pt}$ and $\hat{p}^{\star}_{f2pt}$ respectively. We will compare the pass-through behavior of firm f in destinations 1 and 2 as a "ceteris-paribus" experiment to explore the effect of firm size on pass-through, size being defined as firm f's market share on destination $d = 1, 2$. Let $\hat{P}^{\star}_{1pt}$ and $\hat{P}^{\star}_{2pt}$ be the log-changes in the two destinations' price indices once all adjustments have taken place. Let also λ_1 and λ_2 be short-hand notation for λ_{f1pt} and λ_{f2pt} respectively. Then

$$\begin{aligned} \hat{p}_{f1pt} - \hat{p}_{f2pt} &= (\lambda_1 - \lambda_2)\hat{e}_{ot} + \lambda_1 \hat{P}^{\star}_{1pt} - \lambda_2 \hat{P}^{\star}_{2pt} \\ &= (\lambda_1 - \lambda_2)(\hat{e}_{ot} + \hat{P}^{\star}_{1pt}) + \lambda_2(\hat{P}^{\star}_{1pt} - \hat{P}^{\star}_{2pt}). \end{aligned} \quad (12)$$

As $s^{\star}_{f2pt} < s^{\star}_{f1pt}$, by (4) and (8) it can be shown (see Atkeson and Burstein 2008 or Auer and Schoenle 2016) that $\lambda_1 - \lambda_2 > 0$. Again, $\hat{P}^{\star}_{1pt}$ and $\hat{e}_{ot}$ have opposite signs, but $\hat{e}_{ot} + \hat{P}^{\star}_{dpt} > 0$ for $d = 1, 2$, so the whole first term on the RHS of (12) is positive, contributing to a stronger adjustment of firm f's home-currency price (less pass-through) on market 1. By contrast, the second term is negative, as $\lambda_2 > 0$ and $\hat{P}^{\star}_{1pt} < \hat{P}^{\star}_{2pt} < 0$ if $s^{\star}_{f2pt} < s^{\star}_{f1pt}$. Thus, the general direction of the effect is indeterminate.

If price indices are held constant ($\hat{P}^{\star}_{1pt} = \hat{P}^{\star}_{2pt} = 0$), the second term of (12) vanishes and the expression becomes unambiguously positive. With endogenous price-index adjustment, (12) is positive only in a limited range of firm f's size bounded above by a critical value. To see this, write $\hat{p}(s^{\star}_{fdpt})$ as shorthand for firm f's pass-through on destination d as a function of its size (market share) $s^{\star}_{fdpt}$, and note that by Proposition 1, there exists at least one strictly positive value of $s^{\star}_{fdpt}$ such that $\hat{p}(s^{\star}_{fdpt}) > 0$. Moreover, by (10), $\hat{p}(0) = 0$. As all functions in (9) are continuous, there must exist a critical size on destination d, $\tilde{s}^{\star}_{dpt}$, such that, for any $s^{\star} < \tilde{s}^{\star}_{dpt}$, $s^{\star}_{fdpt} < s^{\star} \implies \hat{p}(s^{\star}_{fdpt}) < \hat{p}(s^{\star})$. If $\hat{p}(s^{\star}_{fdpt})$ is a strictly concave function, $\tilde{s}^{\star}_{dpt}$ is the point at which its derivative is zero; otherwise, it may have several critical values and $\tilde{s}^{\star}_{dpt}$ is the smallest of them. From now on, we will say that firm f is "small" in destination d whenever $s^{\star}_{fdpt} < \tilde{s}^{\star}_{dpt}$.

Thus, provided that home firm f's share of destination market 1 is smaller than the critical value, it does less pass-through on market 1, where it is larger, than on destination market 2, where it is smaller; and we can state:

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

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.

We now derive a corollary of Proposition 2 establishing the link between exchange-rate pass-through and pricing-to-market. Again, we consider a home firm f selling a variety of product

p on two identical destinations $d = 1, 2$ and small on both. However, instead of facing identical exchange-rate shocks on both destinations, now firm f faces a shock $\hat{e}_{o1t} > 0$ on destination 1 but none on destination 2; that is, the currency of country 1 appreciates vis-a-vis firm f's home currency (and that of destination 2). Then

$$\hat{p}_{f1pt} - \hat{p}_{f2pt} = \hat{p}_{f1pt} = \lambda_{f1pt} \left(\hat{e}_{o1t} + \hat{P}^{\star}_{1pt} \right) > 0. \tag{13}$$

Thus, we can state

Corollary 1 *(Pricing-to-market): Starting from a symmetric pricing rule $p_{fdpt} = p_{fd'pt} \, \forall \; d, d'$, a given positive bilateral exchange-rate shock in destination d, $\hat{e}_{odt} > 0$, induces a small firm f to introduce a wedge $p_{fdpt} - p_{fd'pt}$ between its price for sale in destination d and its price for sale in any other destination d'. Moreover, the absolute value of this wedge is increasing in firm f's market share $s^{\star}_{fdpt}$ in destination d.*

We next explore the relationship between destination-market trade policy instruments, tariffs and NTMs, and the extent of firm-level PTM. For that, we again consider "ceteris-paribus" experiments comparing the pass-through behavior of home firm f in two destination markets that vary in their trade policy environment. We first consider the case of a non-tariff measure. Let firm f export a given variety of product p to two destinations $d = 1, 2$ and small on both. Suppose that destination 2 is affected by an NTM forcing all active firms to use a more capital-intensive technology with a higher fixed cost $F_2 > F_1$, but that the two destination markets are otherwise identical; in particular, there is no trade cost or tariff to sell in either destination. By (4), markups are increasing in market share; therefore, the higher fixed cost on destination 2 crowds out small firms, a classic result in heterogeneous-firms models. As destinations 1 and 2 are otherwise identical, it follows that for a firm f with productivity levels above the zero-profit condition (see equation (32) in appendix) but with market shares below the critical value $\tilde{s}^{\star}_{2pt}$, $s^{\star}_{f1pt} < s^{\star}_{f2pt}$ so, by Proposition 2, $\hat{p}_{f1pt} < \hat{p}_{f2pt}$: Firm f does more pricing to market on market 2. Thus, we can state

Proposition 3 *(Non-tariff measures and PTM): For a firm f that is active but small on destination market d, an NTM reduces the degree of exchange-rate pass-through and raises the extent of pricing to market.*

In the empirics, we have no time variation in NTMs; thus, in markets with NTMs, the distribution of firms is truncated with firms with market shares below the zero-profit condition market share, out of the sample. Thus, what we expect to see in the data is the effect identified in Proposition 3.

Consider finally the case of a tariff. Consider two destination markets with a symmetric fixed cost F. Suppose that firm f exporting product variety p to both destinations, faces a tariff in market 2 and none in market 1. Except from the tariff in destination 2 there is no other trade cost and the two destinations are identical. As shown in appendix, in the presence of a fixed cost, a tariff in destination market 2 also has an extensive margin effect, displacing small foreign firms in favor of larger remaining firms both domestic and foreign.[11] This means that for the remaining home firm f in destination market 2, with productivity levels above the zero-profit condition but with market shares below the critical value $\tilde{s}^{\star}_{2pt}$, two effects play against one another: (i) the exit of smaller home competitors raises firm f market share in destination 2, while (ii) the higher level of tariff in destination market 2 reduces firm f market share (see equation (30) in appendix). In net, the effect of a tariff in destination market 2 on home firm f market share is inderminate. Thus, by Proposition 2, we can state

Proposition 4 *(Tariffs and PTM): The effect of tariffs on the degree of exchange-rate pass-through and thus on the extent of pricing to market is inderterminate.*

The effect of tariffs on PTM is thus an empirical question. However, a key difference with non-tariff measures is that tariffs are discriminatory in nature while NTMs are not (WTO's national treatment clause). Tariffs are then more likely to displace foreign firms in favor of domestic ones, implying that the second effect dominates and in net tariffs reduce the market shares of exporters and thus their incentive to engage in PTM. In the empirics, we will verify that this is indeed the case using our large dataset of developing and emerging countries.

3 Data and descriptive statistics

Testing our predictions requires gathering three main types of data: (i) firm-level data on export flows, (ii) macroeconomic data and (iii) trade policy variables. Note that when testing the impact of trade policy on market power, the use of firm-level data is key because it enables us to control for firm characteristics which affect their reaction to exchange rate variations (e.g. marginal costs, size or other firm characteristics such as financial constraints) and can be correlated with trade protection.

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 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).[12] 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 exclude mineral products (chapters HS 25 to 27) and services (chapters HS 98 to 99).[13]

[11] See this from equation (32) in appendix, which gives the zero-profit condition productivity level as an increasing function of the level of import tariff.

[12] 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.

[13] Mineral and primary products are commonly disregarded due to large and sudden fluctuations in international prices and associated terms-of-trade shocks, arbitrarily driving the export performance.

Second, we keep only flows over a thousand USD. Third, for both unit values and export volumes we drop all observations belonging to the top and bottom percentiles in terms of levels and growth rates, percentiles being computed by origin country and sector (HS2). Table 1 gives basic information on final sample size and period by origin country.

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	205,839	34,675	6,525	158	3,119	15.9	9.8
Jordan	2004-11	22,443	3,078	2,066	139	1,135	9.9	5.1
Kenya	2006-11	42,759	7,802	2,921	139	2,296	9.4	15.2
Kuwait	2009-10	4,533	2,275	802	73	939	7.2	23.0
Lebanon	2009-10	30,113	15,060	2,497	132	1,689	11.9	33.2
Mexico	2001-09	458,691	91,194	41,516	156	4,471	9.2	25.0
Morocco	2003-10	125,303	15,694	6,293	153	2,373	8.4	12.4
Rwanda	2006-11	763	147	229	41	117	6.1	3.1
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

For comparability of statistics, product are defined at the six-digit level of the Harmonized System.

The sample is dominated by four countries, Bangladesh, Chile, Mexico and Morocco, in terms of transactions (both total and yearly) and number of firms. In the emprical analysis we will report our results on the entire sample as well as split by origin country. All origin countries have diversified destination portfolios, and the total number of HS6 products exported in one year or another ranges between 117 (Rwanda) and 4,471 (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).

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 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:[14]

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

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

[14] 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 0.07 percent of total trade value.

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 origin-destination-product-year (*odpt*) quadruplet, we compute the bilateral applied tariff. As we are mostly interested in the role played by differences in trade policy across markets, rather than in the effect of variations in trade policy in a given market over time, we also compute the average bilateral applied tariff over the period. This allows us to smooth out missing values without much loss of information. In the empirical analysis we will show that our results are robust to using time-varying bilateral tariffs.

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.[15] The binary data was converted into estimated 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 8 in the appendix shows the (unweighted) average levels of bilateral applied tariffs and NTMs AVEs for the 45 countries covered by both tariff and NTM data. The lowest levels of tariffs are observed in developed countries (e.g. 0.8% in Japan). 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 (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 73,000 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%

[15]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.

to 7% on average.[16] The median firm in our sample exports 2 products and serves 1 destinations against 6 products and 3 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).[17]

Table 2: Descriptive statistics

	Obs.	Mean	S.D.	Q1	Median	Q3
Volume (weigth in kg)	1,034,607	181806	2.67E+06	637	5036	38700
ln volume	1,034,607	8.42	3.02	6.46	8.52	10.56
Unit value (LCU)	1,034,607	113126	2.56E+07	56	355	1466
ln unit value	1,034,607	5.75	2.54	4.03	5.87	7.29
Number of products (firm, t_0)	1,034,607	19.77	38.83	2.00	6.00	19.00
ln number of products	1,034,607	1.95	1.42	0.69	1.79	2.94
Number of destinations (firm, t_0)	1,034,607	10.43	16.20	1.00	4.00	12.00
ln number of destinations	1,034,607	1.48	1.30	0.00	1.39	2.48
Real exchange rate	1,010,335	116.46	420.05	4.49	11.07	53.75
ln real exchange rate	1,010,335	2.20	2.83	1.50	2.40	3.98
GDP (constant 2000 USD)	1,022,768	4.18E+12	5.38E+12	8.55E+10	1.13E+12	1.18E+13
ln GDP	1,022,768	27.21	2.57	25.17	27.75	30.10
Bilateral distance (km)	1,034,607	4912.29	3970.37	1991.24	3369.05	7311.51
ln distance	1,034,607	8.18	0.84	7.60	8.12	8.90
Foreign import tariff	955,458	4.78	12.27	0.00	0.18	6.67
ln (1 + tariff/100)	955,458	0.04	0.08	0.00	0.00	0.06
Non-tariff measure (NTM AVE)	205,566	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.

4 Pass-through across countries and firms

We now turn to an empirical test of the predictions derived from our model in section 2. For each prediction we present the empirical approach, the main results and a series of robustness checks. This section, provides an empirical assessment of our first two propositions: (i) Exchange rate pass-through is incomplete, implying that firms price to market, and (ii) pass-through decreases with firm size, implying that larger firms price more to market.

4.1 Average pass-through

We start by providing estimates of the pass-through parameter on our sample of twelve developing countries and comparing it with existing estimates from industrial and emerging countries. Let us denote by $\ln \text{UV}_{fdpt}$ the log of firm f's producer price for eight-digit product p exported to

[16] These numbers are lower than the average levels of protection displayed in Table 8 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.

[17] In Table 2, the median numbers of (HS6) products and destinations appear respectively as 6 and 4. This reflects multiple counting of multi-product multi-destinations exporters at the level of the unit of observation (firm-destination-product).

destination d in year t, proxied by its FOB unit value and expressed in country o's currency. [18] Let e_{odt} be the average real exchange rate between the origin and destination countries in year t as defined in equation (14). Finally, let $\mathbf{x_{dt}}$ be a set of time-varying destination specific controls, including the destination's GDP. Our baseline empirical specification estimates the effect of changes in the real exchange rate on $\ln \text{UV}_{fdpt}$, within firm-destination-product triplets over time:

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

where $\mathbf{FE}_{fdp}$ and $\mathbf{FE}_{ot}$ are respectively firm×destination×product and origin×year fixed effects. In equation (15), an increase in e_{odt} is a depreciation of the home currency of firm f (the exporter). The coefficient of interest β_1 is the elasticity of home-currency export prices to bilateral exchange rates and maps one-to-one into an exchange rate pass-through elasticity (pass-through is complete when $\beta_1 = 0$ and is zero when $\beta_1 = 1$).[19] Equation (15) is estimated by OLS with robust standard errors clustered at the dyad-product-year level.

A potential issue with specification (15) is that firm-level price changes could be affected by supply-side shocks at the product or firm level, which might in turn correlate with aggregate exchange rate movements. To further ensure that we are indeed identifying the causal impact of exchange rate variations on prices, we take advantage of the high dimensionality of our data and show robustness checks where we replace $\mathbf{FE}_{ot}$ by either origin×product×year fixed effects ($\mathbf{FE}_{opt}$) or firm×product×year $\mathbf{FE}_{fpt}$ fixed effects.

Table 3 reports baseline least-squares estimates of equation (15) for the whole sample (column 1) and split by country or country groups (columns 2 to 7).[20] 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 1.5% 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.[21] Thus, although our elasticities are estimated on a sample of developing countries, we find a degree of exchange rate pass-through which is very much 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 (85% in column 1). The estimated 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 pass-through may lower on aggregate. We test for

[18] Note that as our dataset does not contain information on firm ownership, all firms in our sample are treated as independent entities. Thus in the presence of firm subscripts we omit the origin country subscripts.

[19] Note that β_1 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 (See also Mallick and Marques, 2010). This might explain why the literature using yearly trade customs data typically finds lower estimates of pass-through than those found in papers using direct price data at a higher frequency (e.g. Gopinath and Itskhoki, 2010).

[20] Note that in all tables, the number of observations varies depending on the dimensions of fixed effects included, as observations perfectly predicted by the fixed effects (singletons) are dropped.

[21] 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.*, 2015); 0.05 in Italy (Bernard *et al.*, 2013).

the relationship between pass-through and firms heterogeneity in the next subsection.

The second and perhaps more surprising result is that the degree of PTM is very homogenous across origin countries. In all cases but Mexico the elasticity of producer prices to the exchange rate lies between 0.08 and 0.15 (and in robustness checks, the elasticity drops to 0.1-0.2 in Mexico as well – see below). 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 the distribution of firms is skewed, if firms react heterogeneously to exchange rates, one would mechanically expect to find low elasticities in firm-level estimations as small firms, which adjust less at the price margin, represent the majority of observations, driving down estimates compared to those obtained on product-level data.

Table 3: Exchange rate pass-through: Baseline estimates

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dep. var: ln(unit value)							
Exporting countries	All	Excl. Mexico	EAC	MENA	Bangladesh	Chile	Mexico
ln(RER)	0.153^a	0.128^a	0.124	0.088^b	0.126^a	0.129^a	0.282^a
	(0.015)	(0.015)	(0.082)	(0.037)	(0.034)	(0.017)	(0.050)
ln(dest. GDP)	0.052^b	-0.014	-0.193^c	-0.070	0.021	0.064^c	0.489^a
	(0.025)	(0.026)	(0.102)	(0.047)	(0.071)	(0.033)	(0.075)
Observations	670057	425292	38044	133296	91707	162245	244765
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Origin×year FE	Yes	Yes	Yes	Yes	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. A product is defined at the 8 digit level. Observations perfectly predicted by the fixed effects are dropped.

Robustness. Table 9 in the appendix contains a number of robustness exercises. In Panels A and B we add origin×product×year and firm×product×year fixed effects respectively. Although some of the point estimates become statistically insignificant—which is not surprising given the important loss of power in these regressions—the results are globally similar to our baseline results in Table 3. Pass-through is large but less than complete, and its magnitude is quite stable across countries. Finally, we also estimate the equivalent of equation (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 pass-through 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 incomplete pass-through should raise demand in the destination market and firm-level export volumes. We find positive and significant but small elasticities to the exchange rate (Panel C of Table 9). Differences across origin countries are slightly larger, but the coefficients are also less precisely estimated.

4.2 Heterogeneous pass-through across firms

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 power. 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.

While very large, our dataset 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 on PTM. 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. More seriously, variations in firm performance and prices may be simultaneously affected by omitted variables. We address these problems in two ways. First, we measure our size proxies (product scope and destinations served) at the broader firm level rather than at the firm-destination-product at which regressions are run. Second, we use beginning-of-period values.

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

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

where β_1 measures the average exchange rate elasticity of unit values and β_2 the 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.

Table 4 columns (1) and (2) 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 or the number of destinations, all taken at the beginning of each firm's sample period. Consistent with Berman *et al.* (2012) and Amiti *et al.* (2014), among others, the results clearly suggest that large exporters adjust more their prices than small ones to exchange rate movements. Using the number of products or the number of destinations served by the firm as a proxy for its size makes little difference. 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 0.5%, while a firm selling ten products will raise its price by 1.6%—three times more.

Table 4: Firm heterogeneity and exchange rate pass-through

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. var.			ln unit value			
ln(RER)	0.055^b	0.076^b			0.042	0.011
	(0.027)	(0.032)			(0.028)	(0.041)
ln(dest. GDP)	0.035	0.034				
	(0.026)	(0.026)				
ln(RER) × ln(# prod_{t0})	0.048^a		0.057^a		0.026^c	
	(0.011)		(0.014)		(0.014)	
ln(RER) × ln(# dest_{t0})		0.034^a		0.081^a		0.032^b
		(0.012)		(0.015)		(0.014)
Observations	631348	631348	477355	477355	343290	343290
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes
Origin×year FE	Yes	Yes	Yes	No	No	No
Dyad×product×year FE	No	No	Yes	Yes	Yes	Yes
Firm×product×year FE	No	No	No	No	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}) and ln(# dest_{t0}) are the number of products and the number of destinations of the firm during the first year it appears in the sample. A product is defined at the 8 digit level.

Robustness. A potential issue with equation (16) is that the heterogeneity in PTM picked up by the coefficient β_2 could be driven by 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 PTM).

To ensure that we indeed capture firm-performance effects, we again take advantage of the data's high dimensionality to 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:

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

In this demanding specification, β_2 captures the differences in pass-through elasticties 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. Thus, we unambiguously identify the effect of firm characteristics on PTM, as all heterogeneity in PTM across products or destinations is controlled for by $\mathbf{FE}_{odpt}$. 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 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 pass-through elasticity). Columns (3) and (4) of Table 4 show that the result—that larger exporters price more to market—is very robust; if anything, it is reinforced quantitatively.[22]

In columns (5) and (6), we directly control for all supply-side determinants of prices (e.g. productivity) by including the appropriate set of fixed effects. The coefficient on the interaction term between firm size and exchange rate remains positive and statistically significant.[23]

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

We now turn to the core contribution of the paper and test our predictions relating trade policy measures and the level of exchange rate pass-through to export prices (Proposition 3 for non-tariff measures and Proposition 4 for tariffs).

5.1 Non-tariff measures and pricing-to-market

As discussed earlier, tariffs and NTMs have starkly different effects on market structure. NTMs have rent-shifting effects that erode the market power of foreign exporters; only if they are discriminatory; if they are not, they generate compliance costs for *all* producers that induce the exit of the smallest ones, raising the market share of all remaining ones, including foreign exporters. In that case, the market power of foreign exporters is enhanced by NTMs which raise their incentive to engage in PTM (Proposition 3). Thus, our approach provides a test of whether NTMs are, on average, applied in discriminatory fashion, an important and unexplored policy issue. Our empirical test for the effect of NTMs on market structure, using data on ad-valorem equivalents at the destination country-product level, is based on the following specification:

$$\ln \mathrm{UV}_{fdpt} = \beta_1 \ln e_{odt} + \beta_2 \left(\ln e_{odt} \times \mathrm{NTM}_{dp}\right) + \gamma \mathbf{x_{dt}} + \mathbf{FE}_{ot} + \mathbf{FE}_{fdp} + \varepsilon_{fdpt} \tag{18}$$

where NTM_{dp} is the ad-valorem equivalent of NTMs imposed by destination d on product p. Unlike tariffs, NTMs are all recorded as "MFN", i.e. applying to all origin countries, a convention which largely reflects the way most of them are administered.[24] In equation (18), $\beta_2 > 0$ is our test for the combined hypothesis that NTMs are applied in non-discriminatory fashion but reduce competition.

Table 5 reports estimates from specification (18). The number of observations is much lower than previously due to the incomplete availability of NTMs AVEs. In spite of the reduced sample size and the attenuation bias due to the fact that AVEs are themselves econometric estimates, the results strikingly confirm the hypothesis of non-discrimination-cum-reduced competition. PTM is significantly stronger quantitatively in markets with high levels of NTMs. Specifically, the

[22] As an alternative, 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.

[23] We have also re-estimated Table 4 taking export volume as the dependent variable. The results clearly show that the exported volumes of high-performance firms react less to exchange rate variations (see Table 10 in appendix). 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) of Table 10 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 1.5%.

[24] 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.

coefficients on the exchange rate and its interaction with the NTM variable in columns (1) and (2) suggest that moving from zero to a 10% ad valorem equivalent raises the price elasticity to exchange rate by a third from 9% to 12%. In column (2) we control for firm performance. In columns (3) and (4), splitting the sample according to the level of NTMs yields similar results: The exchange rate coefficients is twice larger in countries and sectors belonging to the top quartile of NTMs (column (4)) than in those in the first quartile (column (3)). In columns (1)-(4) we cluster standard errors by product×dyad× year; boostrapping standard errors to account for the fact that AVEs are estimated leaves our results unchanged (column (5)). In column (6) we include additional interactions between the exchange rate and country or sector dummies. In columns (7) and (8) we control for firm×product×year fixed effects. The coefficients are smaller in magnitude and only significant in column (7). This is not surprising, given that the specification in column (8) is extremely demanding as we include all the previous controls simultaneously.

Finally, in columns (9) and (10) we further test the mechanisms through which NTMs affect market structure. We re-estimate equation (18) on two different samples, based on the share of the firm in destination-market imports. Import shares are computed as in Table 6 and are evaluated, as before, at the beginning of the period. The coefficient on the interaction term between the exchange rate and NTM AVEs is only positive and significant in column (10), and it is more than three times larger (in absolute value) in the sample of larger firms.

Table 13 in the appendix includes both trade policy measures simultaneously. The results are extremely robust. Similarly, alternative clustering strategies, at the firm or dyad-year levels do not alter the statistical significance of the results.

Table 5: Non tariff measures and exchange rate pass-through

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Dep. var. ln unit values										
Variable			NTM AVE						Import share	
Subsample			Low	High					Low	High
ln(RER)	0.094^a	0.106^a	0.092^a	0.182^a	0.094^a	-2.541	0.100^a	0.289	0.210	0.076^a
	(0.021)	(0.035)	(0.024)	(0.037)	(0.018)	(2.636)	(0.019)	(1.328)	(0.131)	(0.025)
ln(dest. GDP)	0.113^a	0.112^a	0.171^a	-0.057	0.113^a	0.129^a	0.074^b	0.085^b	0.059	0.117^b
	(0.043)	(0.043)	(0.050)	(0.074)	(0.038)	(0.044)	(0.038)	(0.037)	(0.173)	(0.054)
ln RER × NTM AVE	0.255^a	0.256^a			0.255^a	0.161^a	0.095^c	-0.004	-0.112	0.340^a
	(0.050)	(0.050)			(0.037)	(0.060)	(0.056)	(0.064)	(0.373)	(0.051)
ln RER × ln(# product$_{t0}$)		-0.006						-0.012		
		(0.016)						(0.023)		
Observations	134705	134705	98231	34874	192724	134705	76506	76506	7276	68932
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Origin×year FE	Yes	Yes	Yes	Yes	Yes	Yes	No	No	Yes	Yes
Firm×product×year FE	No	No	No	No	No	No	Yes	Yes	No	No
Additional controls[1]	No	No	No	No	No	Yes	No	Yes	No	No

^{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 (2) which shows bootstrapped standard errors. High and Low means respectively above the third quartile or below the first quartile of the corresponding variable. [1]: additional controls include additional interaction terms between the RER variable and origin, destination, and sector (HS2) dummies.

5.2 Tariffs and pricing-to-market

We now turn to the results on tariffs. Proposition 4 implies an ambiguous effect of tariffs on foreign firms market shares and thus on their incentives to engage in PTM. How import tariffs in one destination market affect foreign firms market power is thus an emripirical questions which we take to the data. We use data on applied tariffs which vary across product-destination cells and take into account all preferential regimes. In order to make the exercise comparable to the similar one we run on non-tariff measures, for which we have no time-wise variation, we use a time-wise average of tariffs over the sample period.[25] The corresponding estimating equation is:

$$\ln \text{UV}_{fdpt} = \beta_1 \ln e_{odt} + \beta_2 \left[\ln e_{odt} \times \ln(1 + t_{odp})\right] + \gamma \mathbf{x_{dt}} + \mathbf{FE}_{ot} + \mathbf{FE}_{fdp} + \varepsilon_{fdpt} \qquad (19)$$

where t_{odp} is the average tariff imposed by destination country d on product p imported from origin o over the period. As per Proposition 4, the sign of β_2 is indeterminate. However as previously discussed, since tariffs are discriminatory we would expect their rent-shifting effects to dominate and in net see them erode the market power of foreign exporters. That is, empirically we would expect β_2 to be negative.

Table 6 reports the results. In column (1), the estimating equation is just (19). In column (2), we control for firm f's size by including an interaction term between its number of products and the exchange rate. In both cases, the coefficient on β_2 is negative and highly significant. We find strong support for rent-shifting effects of tariffs, i.e. tariffs reduce the elasticity of home-currency export prices to the exchange rate, implying a higher pass-through.

How large are the effects? Faced with a 10% depreciation of its home currency, firm f selling a product tariff-free in a given destination would raise its home-currency price by 1.9% (column 1). Faced with the same depreciation on a destination with a 10% tariff, it would raise it by only 1.3%, or about 30% less. When the tariff reaches 30-35% , pass-through is complete, i.e. no significant PTM is detected anymore. Similar results are obtained in columns (3) and (4), which report results for destination-product (dp) cells in the highest quartile of tariffs vs. the lowest (where they are zero). In the latter, firms raise their home-currency price by 2% following a 10% depreciation. In the former, they leave it largely unchanged, implying quasi-complete pass-through. Thus, tariffs at their current levels have substantial effects on the pricing behaviour of foreign exporters, suggesting that they affect market structure (presumably through rent-shifting effects) in a non-trivial way.

Robustness. We next show that our results are robust to the inclusion of additional controls and dimensions of fixed effects. In column (5) we include additional interaction terms between the exchange rate and (i) destination group dummies; (ii) origin country dummies; and (iii) product dummies. Results are robust, suggesting that we are identifying the effect of market-specific trade policy, rather than the role of other country- or product-specific determinants of exchange rate pass-through. In column (6) we include firm×product×year fixed effects. In this specification, we unambiguously identify PTM, as we exploit variations across destination markets, for a given firm-product, and explicitly account for all changes in marginal costs. Finally, column (7) reports results when we include all the previous controls simultaneously.

[25] An additional reason for doing this is that tariff data contain many missing values; using average values allows maximizing the number of observations. We however provide robustness exercises using time-varying tariffs.

In columns (8) and (9), although we do not have market-share data (which is why we identify market power indirectly through pricing behavior), we attempt to proxy, albeit very imperfectly, the size of firms through their share of aggregate imports. Import shares are computed over destination d's total imports of product p, obtained from the BACI database, with products defined at the 4-digit level.[26] These cannot be interpreted as market shares since the sales of domestic firms, which are unobserved, are excluded from the denominator. Therefore, they cannot be used to assess market power. However, comparing two exporters f and f' in terms of their shares of the destination's imports, if f has a lower share, it is closer than f' to our atomistic assumption. As expected, firms with low import shares are the ones for which tariffs have the strongest impact, i.e. for which β_2 is largest in absolute value.

Table 6: Import tariffs and exchange rate pass-through

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Dep. var. ln unit values									
Variable			Tariffs					Import share	
Subsample			Low	High				Low	High
ln(RER)	0.190^a	0.093^a	0.206^a	0.039^c	-2.812	0.104^a	-4.134^a	0.473^a	0.135^a
	(0.019)	(0.030)	(0.034)	(0.023)	(1.901)	(0.017)	(0.893)	(0.109)	(0.021)
ln(dest. GDP)	0.045^c	0.047^c	-0.071	0.250^a	0.043	0.080^a	0.079^a	-0.010	0.121^a
	(0.027)	(0.027)	(0.055)	(0.039)	(0.027)	(0.026)	(0.026)	(0.130)	(0.033)
ln RER × ln(1 +tariff/100)	-0.659^a	-0.641^a			-0.399^b	-0.455^a	-0.255^c	-2.318^a	-0.455^a
	(0.146)	(0.147)			(0.155)	(0.127)	(0.133)	(0.806)	(0.162)
ln RER × ln(# prod$_{t0}$)		0.048^a					0.018		
		(0.011)					(0.015)		
Observations	624329	624329	297324	160828	624329	333045	333045	124929	183033
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Origin×year FE	Yes	Yes	Yes	Yes	Yes	No	No	Yes	Yes
Firm×product×year FE	No	No	No	No	No	Yes	Yes	No	No
Additional controls[1]	No	No	No	No	Yes	No	Yes	No	No

[c] significant at 10%; [b] significant at 5%; [a] significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses.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. [1]: additional controls include additional interaction terms between the RER variable and origin, destination, and sector (HS2) dummies.

Tables 11 and 12 in the appendix report additional robustness checks, using alternative ways of measuring bilateral tariffs. Table 11 replicates Table 6 using time-varying bilateral tariffs instead of period averages. This results in a significant loss of observations, but the results are unaffected. If anything, they are reinforced quantitatively. Note that the coefficient on tariffs alone can now be identified. Interestingly, we find that it is negative, significant and of a similar order of magnitude as the exchange rate coefficient in our baseline estimations (column 2 of Table 3). In other words, tariffs and exchange rate pass-through are found to be of similar (high) levels in our sample. This result is consistent with the findings of Feenstra (1989). In Table 12 we use the level of tariffs instead of its log transformation. The results are again similar. Finally, these results are robust to alternative clustering strategies, at the firm or dyad-year levels.[27]

[26] See http://www.cepii.fr/CEPII/en/bdd_modele/presentation.asp?id=1.

[27] Results available upon request.

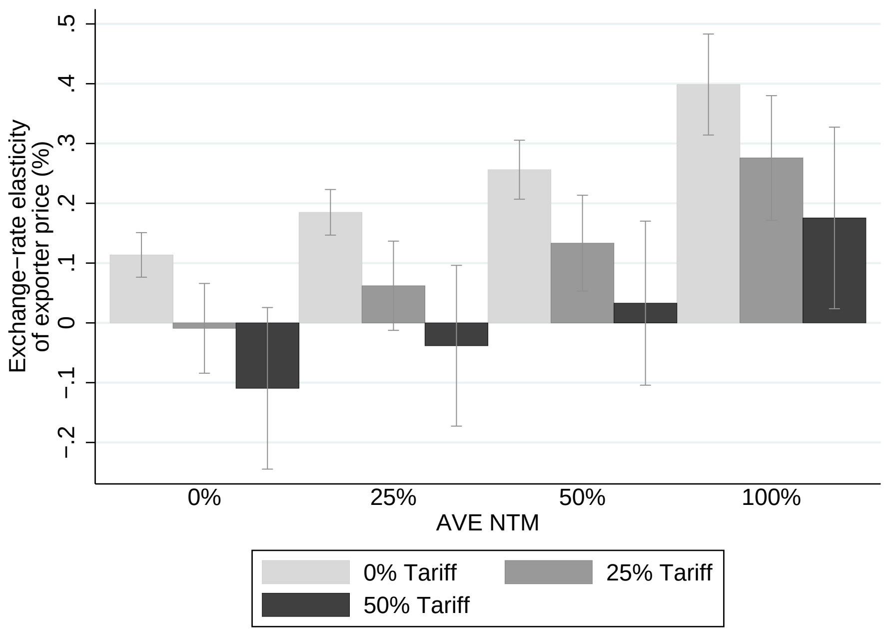
5.3 Differential impact of trade policy on market power

To illustrate the differential effect of tariffs and NTMs on PTM, we take the coefficients from column (1) of Table 13 in the appendix and compute the predicted degree of pass-through for various combinations of tariffs and non tariff measures.

The results are shown in Figure 1. In markets with low-impact NTMs (first three bars on the left), exchange rate pass-through is almost complete, except in tariff-free markets where it declines slightly to around 90% (1-0.1). Pass-through becomes strongly incomplete and shrinks to as low as 60% ($1 - 0.4$) in markets with zero tariffs but high-impact NTMs (last three bars on the right). Again, given the various controls included in our estimations, it is unlikely that we are capturing the effect of other determinants of PTM than trade policy and its impact on market power.

Taken together, these results suggest that, instead of displacing foreign firms in favor of domestic ones like tariffs, NTMs displace small firms in favor of larger ones, including both domestic firms and foreign exporters. Such an effect is consistent with a mechanism whereby NTMs raise compliance costs, inducing the exit of smaller firms, while leaving the remaining (larger) ones with expanded market shares. The notion that NTMs may affect market structure in terms of large vs. small firms more than home vs. foreign is quite new, as most of the literature on NTMs has assumed so far that they were mere surrogates for tariffs (see e.g. Feenstra 1984, Deardorff and Stern 1997, Baldwin 1989, Leamer 1990, Anderson and Neary 1994, Kee, Nicita and Olarreaga 2009, Carrere and de Melo 2011, or Cadot and Gourdon 2014a).

Figure 1: Trade policy and exchange rate pass-through



Predicted effects and confidence bands computed from column (1) of Table 13.

5.4 A further look at import shares

As discussed, import shares are poor proxies for true market shares given that variations in the sales of domestic firms, which are key for the measurement of rent-shifting effects, are not observed. In this section, we try to go around the problem by instrumenting import shares with tariffs and NTMs. That is, our postulate is that the variation of import shares explained by the variation of tariffs and NTMs has something to do with the effect of trade policy on market structure. While this identification strategy is quite indirect, we use it as a further robustness check on our claim that changes in the pricing behavior of exporters reflects the effect of destination trade policy on market structure.

This exercise has a number of limitations, however. First, we cannot directly instrument a firm's import share with trade-policy instruments because it is absorbed by firm-destination-product fixed effects—only the interaction with exchange rates 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 import shares. Third, it is possible that our firm-level customs data and UN-COMTRADE data do not perfectly match, making the computation of import shares problematic. Finally, if tariffs and quotas were affecting home-currency producer prices through channels other than market power, they would not be valid instruments; however, this is unlikely given that our prices are demeaned in all key dimensions through fixed effects (for instance, large-country tariff effects on world prices would be washed away by demeaning across destinations). For all these reasons, we consider this exercise as a complement to our baseline results shown in Table 6 and Table 5.

Table 7 column (1) reports the results of a preliminary analysis where we include in our baseline specification (equation (15)) an interaction term between the real exchange rate and the firm's average import share in destination-product cell d, p. 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 destination for that HS4 product. The coefficient on the interaction term between import shares and the real exchange rate is positive and statistically significant at the 5 percent level, a result consistent with Amiti *et al.* (2014).

As average import shares are however potentially endogenous (influenced by prices and omitted variables), we instrument their interaction with exchange rate by interaction terms between the exchange rate and our two trade policy measures. We expect firms with larger import shares due to low tariffs or high-impact NTMs to adjust more their prices in response to exchange-rate shocks. This is what we find (see also Table 14 in the appendix).[28]

Columns (2) and (3) of Table 7 use 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 PTM, when instrumented by trade policy. This effect survives when controlling for firm-product-year fixed effects in column (3). Column (4) and (5) use 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

[28]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.

fact that tariffs vary more, as they are *de facto* bilateral, contrary to NTMs.

For all their limitations, these results suggest that our baseline estimates of Tables 6 and 5 reflect, as we claim, the effect of trade policy on market power rather than that of confounding influences.

Table 7: Import share, trade policy and exchange rate pass-through

Dep. var: ln unit value	(1)	(2)	(3)	(4)	(5)
ln(RER)	0.107^a	-0.243	-0.344	-0.239	-0.160
	(0.023)	(0.181)	(0.241)	(0.165)	(0.119)
ln(dest. GDP)	0.152^a	0.227^a	0.151^b	0.199^a	0.088^c
	(0.046)	(0.075)	(0.069)	(0.073)	(0.050)
ln (RER) × import share$_{fdp}$	0.401^b	11.132^b	11.810^b		
	(0.199)	(5.418)	(5.964)		
ln (RER) × import share$_{odp}$				1.838^b	1.223^b
				(0.773)	(0.522)
First stage (dep. var.: ln RER × import share)					
ln (RER) × ln(tariff+1)	-0.066^a	-0.055^a	-0.507^a	-0.553^a	
	(0.016)	(0.020)	(0.151)	(0.197)	
ln (RER) × NTM AVE	0.002	0.003	0.098	0.083	
	(0.005)	(0.006)	(0.101)	(0.116)	
ln(RER)	0.035^a	0.039^a	0.204^a	0.224^a	
	(0.003)	(0.004)	(0.027)	(0.034)	
ln(dest. GDP)	-0.006^a	-0.005^c	-0.018^a	0.003	
	(0.002)	(0.003)	(0.007)	(0.013)	
Observations	123135	123135	67354	121445	66877
Hansen overid p-value	-	0.02	0.14	0.43	0.74
F-stat excl. instruments	-	8.3	4.6	6.1	6.6
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes
Origin×year FE	Yes	Yes	No	Yes	No
Firm×product×year FE	No	No	Yes	No	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.

6 Concluding remarks

Our objective in this paper is to provide large-scale evidence that trade policy has strong effects on market power at the firm-level and that the direction of this effect varies with the type of instruments used. The question is not new. Yet, existing papers typically focus on specific countries, industries and trade policy instruments, essentially because identifying indicators of market power—such as price-cost margins or markets shares—on a wide sample of countries and sectors is problematic. The novelty of our paper is to circumvent this issue by borrowing our market-power proxy from the PTM literature, identifying it indirectly by observing the pricing behavior of exporting firms on their different destination markets in reaction to bilateral exchange rate shocks. As destination markets vary in terms of their trade protection, we can then infer the effect of trade policy on market power; moreover, as our identification relies on an interaction term (between bilateral exchange rate shocks and destination-specific trade-policy measures) we can control for a large set of confounding influences at the level of both the firm and the market with a powerful array of fixed effects. All we need is detailed data on firm-level export prices.

We find that PTM is prevalent, although quantitatively limited, in all regions of our sample. 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 is surprisingly little variation in this split, even though our sample spans several continents and included countries at different levels of development and integration in global value chains. We also find, on the basis of various proxies for firm size (and hence performance), that PTM clearly rises with exporter size, in accordance with existing evidence. Yet, the prevalence of PTM in a developing-country sample consisting predominantly of firms that are, by international standards, small to medium-sized, comes as a surprise as PTM is typically associated with relatively large firms.

More importantly, we provide robust evidence that trade policy deeply affects market structure. In accordance with theory and intuition, exporters faced with tariffs on their destination markets do significantly less PTM, revealing a loss of market power consistent with rent-shifting effects.

Last but not least, our aproach sheds new light on the effect of NTMs. While trade economists have typically treated NTMs as surrogates for tariffs and focused on their trade-inhibiting or rent-shifting effects, we show that they affect market structure quite differently from tariffs. High-impact NTMs reduce the degree of pass-through of foreign exporters, as if they had *more* market power. Although we cannot measure directly firm entry and exit into NTM-ridden markets for lack of time-wise variation in NTMs, our findings are consistent with the following conjecture: NTMs today are largely regulatory interventions (sanitary, technical, etc.). Those interventions require firms to adapt their production technology, which may crowd out the least efficient ones, whether domestic or foreign. More efficient ones, again irrespective of whether they are domestic or foreign, benefit from this change in market structure with expanded market shares. We detect this through their pass-through/PTM behavior, and the effect seems extremely robust to a variety of robustness tests.

This result is quite new and suggests that, from a policy standpoint, NTMs (by which we mean measures such as technical or sanitary regulations) might well be viewed as a competition-policy issue as much as a trade one. This has implications for policy. For instance, as part of deep-integration schemes under ASEAN and the proposed Trans-Pacific Partnership (TPP), a number of countries in Asia are considering setting up regulatory supervision bodies focused on NTMs. Our approach suggests that the mandate of such bodies should encompass the impact of NTMs on domestic market structure rather than just on trade facilitation. This might be best achieved by merging them with existing competition commissions or giving them authority over both regulatory supervision *and* competition policy.

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Appendix 1: The model

In order to keep the notation in this appendix consistent with that in the empirical part of the paper but nevertheless reasonably light, p as a variable is a price while p as a subscript is a product.

Technology

Production takes place under constant returns to scale. Let a_{fp} be firm f's technology, drawn from a distribution that we do not need to specify, and ℓ_{fdpt} the labor it allocates to the production of product p for sale in destination d. Output is

$$q_{fdpt} = a_{fp}\ell_{fdpt}. \tag{20}$$

Let w_{ot} be the wage rate in origin country o at t. Firm f's marginal cost is $c_{fpt} = w_{ot}/a_f$ in home currency and $c^{\star}_{fdpt} = w_{ot}/(a_{fp}e_{odt})$ in foreign currency.

Preferences

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 finite number of domestic and foreign firms, indexed by f. The elasticity of substitution is η between products and ρ between varieties, with $1 < \eta < \rho$.

The upper-level aggregate, final consumption, is:

$$Q_{dt} = \left[\int_p (Q_{dpt})^{\frac{\eta-1}{\eta}}\right]^{\frac{\eta}{\eta-1}} \tag{21}$$

with price index

$$P^{\star}_{dt} = \left[\int_p (P^{\star}_{dpt})^{1-\eta}\right]^{\frac{1}{1-\eta}}. \tag{22}$$

Going down one level, the inverse demand function for product p is

$$\frac{P^{\star}_{dpt}}{P^{\star}_{dt}} = \left(\frac{Q_{dpt}}{Q_{dt}}\right)^{\frac{-1}{\eta}} \tag{23}$$

The CES aggregator of firm-level varieties into products (the "lower-level") is:

$$Q_{dpt} = \left[\sum_f (q_{fdpt})^{\frac{\rho-1}{\rho}}\right]^{\frac{\rho}{\rho-1}}. \tag{24}$$

Pricing

The only trade cost is an ad-valorem tariff $\tau = 1 + t$. Thus, consumers in the foreign country face price $\tau p^{\star}_{fdpt}$, whereas the firm receives $p^{\star}_{fdpt}$. Let $\phi_{fdpt} \in \{0; 1\}$ be equal to one when firm f decides to sell its variety of product p in market d given its productivity draw a_{fp} and the cost of selling in destination d. Let $n(d)$ be the number of firms for whom destination d is their home market, and $n^\star(d)$ the number of firms exporting to destination d from other countries. In the presence of a trade cost τ, the price index in destination d is

$$P^\star_{dpt} = \left[\sum_{i=1}^{n(d)} \phi_{fdpt} {p^\star_{fdpt}}^{1-\rho} + \tau^{1-\rho} \sum_{i=1}^{n^\star(d)} \phi_{fdpt} {p^\star_{fdpt}}^{1-\rho} \right]^{\frac{1}{1-\rho}}. \tag{25}$$

If firm f decides to sell on market d, using (21)-(25), it faces a direct demand function given by

$$q_{fdpt} = \left(\frac{\tau p^\star_{fdpt}}{P^\star_{dpt}} \right)^{-\rho} \left(\frac{P^\star_{dpt}}{P^\star_{dt}} \right)^{-\eta} Q_{dt}. \tag{26}$$

Firm f's profit-maximization problem, given a fixed cost F_{dp} to sell any variety of product p in destination d and demand as in (26) is:

$$\max_{q_{fdpt}, \phi_{fdpt}} \phi_{fdpt} \left(p^\star_{fdpt} q_{fdpt} - q_{fp} c^\star_{fpt} - F_{dp} \right) \quad \text{s.t.} \quad q_{fdpt} = \left(\frac{\tau p^\star_{fdpt}}{P^\star_{dpt}} \right)^{-\rho} \left(\frac{P^\star_{dpt}}{P^\star_{dt}} \right)^{-\eta} Q_{dt} \tag{27}$$

and its optimal pricing rule is

$$p^\star_{fdpt} = \mu_{fdpt}\, c^\star_{fpt} = \left(\frac{\varepsilon_{fdpt}}{\varepsilon_{fdpt} - 1} \right) c^\star_{fpt} \tag{28}$$

where μ_{fdpt} is the markup and

$$\varepsilon_{fdpt} = \frac{\rho\eta}{\rho s^\star_{fdpt} + \eta(1 - s^\star_{fdpt})} > 1 \tag{29}$$

is the price elasticity of demand perceived by firm f. Unlike in standard CES models, μ_{fdpt} is not constant; instead, it varies with firm f's market share. Let us define this market share in nominal foreign-currency terms:

$$s^\star_{fdpt} = \frac{p^\star_{fdpt}}{P^\star_{dpt}} \frac{q_{fdpt}}{Q_{dpt}} = \frac{\tau p^\star_{fdpt}\, q_{fdpt}}{\sum_{i=1}^{n(d)} \phi_{fdpt} p^\star_{fdpt} q_{fdpt} + \tau \sum_{i=1}^{n^\star(d)} \phi_{fdpt} p^\star_{fdpt} q_{fdpt}}. \tag{30}$$

It can be shown that $s^\star_{fdpt}$ is decreasing in τ (the trade cost) and in $n(d)$ and $n^\star(d)$ (respectively the number of domestic firms and exporters selling varieties of product p on destination market d). The elasticity of the markup to the market share is

$$\begin{aligned} \Gamma_{fdpt} &= \left. \frac{\partial \ln \mu_{fdpt}}{\partial \ln \varepsilon_{fdpt}} \frac{\partial \ln \varepsilon_{fpdt}}{\partial \ln s^\star_{fdpt}} \right|_{P^\star_{dpt} \text{constant}} = \left(\frac{1}{1 - \varepsilon_{fdpt}} \right) \frac{\rho - \eta}{\rho\eta} s^\star_{fdpt} \\ &= \left\{ \frac{s^\star_{fdpt}}{1 - [(1 - s^\star_{fdpt})/\rho] - s^\star_{fdpt}/\eta} \right\} \left(\frac{1}{\eta} - \frac{1}{\rho} \right); \end{aligned} \tag{31}$$

as shown by Edmond et al. (2015), the *inverse* markup is a linear function of the market share.

Entry

Firm f variable profit can be rewritten as

$$\pi_{fdpt} = \left(p^{\star}_{fdpt} - c^{\star}_{fpt}\right) q_{fdpt} = (\mu_{fdpt} - 1) \frac{w_{ot} e_{odt} q_{fdpt}}{a_{fp}}. \tag{32}$$

Firm f's zero-profit condition $\pi_{fdpt} = F_{dp}$ determines implicitly a cutoff productivity level $a_{fp}(\tau, F_{dp})$. A rise in F_{dp}, ceteris paribus, sorts firms by their productivity draw a_{fp}, inducing the exit of all firms with productivity below $a_{fp}(\tau, F_{dp})$.[29] This in turn reduces $n(d)$ and $n^{\star}(d)$ and by (30), raises the market share of the remaining firms. As a result, by (4), markups rise. A rise in τ has a similar sorting effect, but, by (30), it also has a negative effect on the market share of exporters selling in d from other countries (the classic rent-shifting effect of tariffs). As a result, its effect on the markups of foreign exporters is ambiguous.

[29] As discussed by Edmond et al. (2015), a Cournot game with endogenous entry has multiple equilibria; we follow them and Atkeson and Burstein (2008) in assuming that firms enter sequentially by decreasing order of a_{fp} and exit also sequentially by increasing order of a_{fp}.

Appendix 2: additional tables

Table 8: Summary statistics: Trade protection

Country	NTM		Applied Tariff		Country	NTM		Applied Tariff	
	mean	s.d.	mean	s.d.		mean	s.d.	mean	s.d.
Argentina	0.16	0.43	0.04	0.04	Lebanon	0.02	0.10	0.04	0.07
Austria	0.11	0.35	0.00	0.00	Lithuania	0.18	0.38	0.01	0.06
Bangladesh	0.17	0.45	0.17	0.11	Luxembourg	0.19	0.46	0.00	0.00
Bolivia	0.01	0.12	0.04	0.04	Madagascar	0.03	0.23	0.03	0.05
Brazil	0.18	0.40	0.04	0.04	Mauritius	0.03	0.30	0.02	0.06
Bulgaria	0.14	0.34	0.00	0.03	Mexico	0.08	0.34	0.03	0.06
Burundi	0.13	0.45	0.05	0.06	Morocco	0.04	0.24	0.03	0.10
Cambodia	0.03	0.24	0.13	0.10	Namibia	0.08	0.41	0.08	0.11
Chile	0.02	0.26	0.01	0.01	Paraguay	0.02	0.18	0.05	0.05
China	0.26	0.52	0.04	0.04	Peru	0.01	0.11	0.06	0.04
Colombia	0.02	0.16	0.06	0.05	Poland	0.14	0.32	0.01	0.06
Czech Republic	0.15	0.35	0.05	0.08	Senegal	0.01	0.12	0.06	0.07
Ecuador	0.05	0.25	0.06	0.06	Slovak Republic	0.16	0.37	0.05	0.08
Egypt, Arab Rep.	0.17	0.57	0.08	0.70	Slovenia	0.18	0.37	0.01	0.04
Finland	0.18	0.37	0.00	0.00	South Africa	0.05	0.23	0.08	0.11
Hungary	0.03	0.32	0.02	0.05	Sri Lanka	0.19	0.50	0.04	0.08
India	0.19	0.52	0.02	0.05	Syrian Arab Republic	-0.01	0.27	0.15	0.18
Indonesia	0.07	0.32	0.02	0.04	Tanzania	0.00	0.11	0.03	0.04
Japan	0.10	0.34	0.01	0.03	Tunisia	0.05	0.24	0.11	0.24
Kazakhstan	0.15	0.40	0.05	0.04	Uganda	0.20	0.46	0.05	0.06
Kenya	0.19	0.47	0.04	0.06	Uruguay	0.00	0.10	0.04	0.04
Lao PDR	0.19	0.45	0.08	0.06	Venezuela, RB	0.07	0.29	0.06	0.06
Latvia	0.17	0.37	0.04	0.07	**Average**	0.10	0.32	0.05	0.08

Source: see section 3.3.

Table 9: Exchange rate pass-through across countries: Robustness

Exporting countries	(1) All	(2) Excl. Mexico	(3) EAC	(4) MENA	(5) Bangladesh	(6) Chile	(7) Mexico
A. Unit values, origin×product×year FE							
ln(RER)	0.123[a]	0.114[a]	0.069	0.114[a]	0.055[b]	0.136[a]	0.181[a]
	(0.012)	(0.012)	(0.089)	(0.035)	(0.025)	(0.013)	(0.036)
ln(dest. GDP)	0.109[a]	0.068[a]	0.151	0.082[c]	0.085[c]	0.042[c]	0.345[a]
	(0.022)	(0.022)	(0.124)	(0.050)	(0.051)	(0.024)	(0.075)
Observations	644699	406715	34009	125510	90555	156641	237984
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Origin×product×year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
B. Unit values, firm×product×year FE							
ln(RER)	0.087[a]	0.084[a]	0.061	-0.023	0.033	0.119[a]	0.080[b]
	(0.013)	(0.013)	(0.108)	(0.044)	(0.031)	(0.013)	(0.040)
ln(dest. GDP)	0.077[a]	0.107[a]	0.244[c]	0.139[b]	0.120[b]	0.073[a]	-0.092
	(0.023)	(0.023)	(0.133)	(0.061)	(0.056)	(0.024)	(0.080)
Observations	368067	273430	22330	63994	62648	124458	94637
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm×product×year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
C. Export volumes, firm×product×year FE							
ln(RER)	0.217[a]	0.272[a]	-0.081	0.405[a]	0.288[b]	0.245[a]	0.088
	(0.042)	(0.047)	(0.182)	(0.117)	(0.144)	(0.057)	(0.087)
ln(dest. GDP)	1.331[a]	1.164[a]	0.168	0.660[a]	1.779[a]	1.362[a]	2.173[a]
	(0.078)	(0.086)	(0.298)	(0.155)	(0.258)	(0.114)	(0.180)
Observations	368067	273430	22330	63994	62648	124458	94637
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm×product×year FE	Yes	Yes	Yes	Yes	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. A product is defined at the 8 digit level.

Table 10: Firm heterogeneity and exchange rate pass-through: Export volumes

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. var.	ln export volumes					
ln(RER)	0.509^a	0.732^a			0.286^a	0.419^a
	(0.063)	(0.072)			(0.080)	(0.110)
ln(dest. GDP)	1.353^a	1.351^a				
	(0.065)	(0.066)				
ln(RER) × ln(# $prod_{t0}$)	-0.159^a		-0.264^a		0.040	
	(0.024)		(0.037)		(0.033)	
ln(RER) × ln(# $dest_{t0}$)		-0.243^a		-0.369^a		-0.020
		(0.027)		(0.041)		(0.040)
Observations	631348	631348	477355	477355	343290	343290
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes
Origin×year FE	Yes	Yes	Yes	No	No	No
Dyad×product×year FE	No	No	No	Yes	Yes	Yes
Firm×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(# $prod_{t0}$) and ln(# $dest_{t0}$) are the number of products and the number of destinations of the firm during the first year it appears in the sample. A product is defined at the 8 digit level.

Table 11: Import tariffs and exchange rate pass-through: Time varying tariffs

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Dep. var. ln unit values									
Variable			Tariffs					Import share	
Subsample			Low	High				Low	High
ln(RER)	0.212^a	0.098^a	0.237^a	0.014	3.552	0.120^a	-2.321	0.511^a	0.150^a
	(0.023)	(0.035)	(0.032)	(0.026)	(4.270)	(0.020)	(3.955)	(0.115)	(0.025)
ln(dest. GDP)	0.018	0.020	-0.223^a	0.215^a	-0.006	0.086^a	0.083^a	0.034	0.098^b
	(0.033)	(0.033)	(0.062)	(0.052)	(0.033)	(0.030)	(0.030)	(0.143)	(0.042)
ln($tariff_t$+1)	-0.063	-0.066^c			-0.052	-0.073^c	-0.066^c	0.870^c	-0.111^a
	(0.039)	(0.039)			(0.041)	(0.039)	(0.040)	(0.493)	(0.042)
ln (RER) × ln($tariff_t$+1)	-1.209^a	-1.204^a			-0.519^a	-0.508^a	-0.182	-2.467^a	-0.887^a
	(0.189)	(0.189)			(0.195)	(0.167)	(0.180)	(0.880)	(0.229)
ln (RER) × ln(# $product_{t0}$)		0.055^a					0.034^b		
		(0.012)					(0.017)		
Observations	540418	540418	327480	134844	540418	265784	265784	118650	144296
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Origin×year FE	Yes	Yes	Yes	Yes	Yes	No	No	Yes	Yes
Firm×product×year FE	No	No	No	No	No	Yes	Yes	No	No
Additional controls[1]	No	No	No	No	Yes	No	Yes	No	No

[c] significant at 10%; [b] significant at 5%; [a] significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses. 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. [1]: additional controls include additional interaction terms between the RER variable and origin, destination, and sector (HS2) dummies.

Table 12: Import tariffs and exchange rate pass-through: Tariff level

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Dep. var. ln unit values									
Variable			Tariffs					Import share	
Subsample			Low	High				Low	High
ln(RER)	0.171^a	0.073^b	0.206^a	0.039^c	-2.822	0.090^a	-4.156^a	0.469^a	0.124^a
	(0.017)	(0.029)	(0.034)	(0.023)	(1.900)	(0.015)	(0.895)	(0.107)	(0.020)
ln(dest. GDP)	0.043	0.045^c	-0.071	0.250^a	0.042	0.079^a	0.079^a	-0.011	0.120^a
	(0.027)	(0.027)	(0.055)	(0.039)	(0.027)	(0.026)	(0.026)	(0.130)	(0.033)
ln (RER) × tariff	-0.272^a	-0.260^a			-0.123^b	-0.181^a	-0.091^c	-2.099^a	-0.221^a
	(0.068)	(0.068)			(0.061)	(0.055)	(0.052)	(0.717)	(0.074)
ln (RER) × ln(# product_{t0})		0.049^a					0.018		
		(0.011)					(0.015)		
Observations	624329	624329	297324	160828	624329	333045	333045	124929	183033
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Origin×year FE	Yes	Yes	Yes	Yes	Yes	No	No	Yes	Yes
Firm×product×year FE	No	No	No	No	No	Yes	Yes	No	No
Additional controls[1]	No	No	No	No	Yes	No	Yes	No	No

[c] significant at 10%; [b] significant at 5%; [a] significant at 1%. Robust standard errors clustered by product-origin-destination-year are in parentheses.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. [1]: additional controls include additional interaction terms between the RER variable and origin, destination, and sector (HS2) dummies.

Table 13: Non tariff measures, tariffs and exchange rate pass-through

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dep. var. ln unit values								
Variable			NTM AVE					
Subsample			Low	High				
ln(RER)	0.114^a	0.114^a	0.092^a	0.182^a	0.125^a	-6.792	0.108^a	0.793
	(0.023)	(0.023)	(0.024)	(0.037)	(0.038)	(4.432)	(0.022)	(2.073)
ln(dest. GDP)	0.143^a	0.143^a	0.171^a	-0.057	0.143^a	0.150^a	0.090^b	0.102^b
	(0.046)	(0.042)	(0.050)	(0.074)	(0.046)	(0.047)	(0.041)	(0.041)
ln (RER) × ln(tariff+1)	-0.550^b	-0.550^b			-0.554^b	-0.087	-0.530^a	0.003
	(0.215)	(0.264)			(0.215)	(0.245)	(0.205)	(0.260)
ln (RER) × NTM AVE	0.285^a	0.285^a			0.285^a	0.178^a	0.133^b	0.036
	(0.052)	(0.049)			(0.052)	(0.063)	(0.059)	(0.067)
ln (RER) × ln(# product_{t0})					-0.005			-0.013
					(0.016)			(0.024)
Observations	127686	183356	98231	34874	127686	127686	71369	71369
Firm×product×dest. FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Origin×year FE	Yes	Yes	Yes	Yes	Yes	Yes	No	No
Firm×product×year FE	No	No	No	No	No	No	Yes	Yes
Additional controls[1]	No	No	No	No	No	Yes	No	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, except in column (2) which shows bootstrapped standard errors. High and Low means respectively above the third quartile or below the first quartile of the corresponding variable. [1]: additional controls include additional interaction terms between the RER variable and origin, destination, and sector (HS2) dummies.

Table 14: Import share, trade policy and exchange rate pass-through

Dep. var: Import share$_{fdpt}$	(1)	(2)	(3)
ln(1 + tariff)	-0.034^a	-0.009^a	-0.009^b
	(0.004)	(0.003)	(0.004)
NTM AVE	0.005^a	0.006^a	0.006^a
	(0.001)	(0.002)	(0.002)
Observations	96388	95988	95946
Adj. R^2	0.050	0.295	0.304
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.

"Sur quoi la fondera-t-il l'économie du monde qu'il veut gouverner? Sera-ce sur le caprice de chaque particulier? Quelle confusion! Sera-ce sur la justice? Il l'ignore."

Pascal

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