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JEL Classification: F10, F13, F14

Keywords: WTO, import tariff, trade policy, exports

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Trade Liberalization and the Extensive Margin of Differentiated Goods: Evidence from China*

Johannes Van Biesebroeck§ Yingting Yi† Elena Zaurino‡

March 2021

Abstract

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1. Introduction

The controversial finding of Rose (2004)—a lack of evidence that the WTO has increased world trade[1]—kicked off a lively debate to assess various dimensions of the WTO's effectiveness to promote international trade. Focusing on sector-level bilateral trade volumes in the gravity framework, Subramanian and Wei (2007) find that the effect on trade growth is strong and robust, but highly uneven across countries and sectors. Distinguishing between the number of products traded (extensive margin) and the average trade volume per product (intensive margin), Dutt, Mihov, and Van Zandt (2013) find that the impact is almost exclusively on the extensive margin. Exploiting variation in tariff rates, rather than relying on a WTO membership dummy, Debaere and Mostashari (2010) find only weak effects of tariff reductions on the number of imported products in the United States. Using export data for French firms, Buono and Lalanne (2012) further show that the extensive margin effects are the result of existing exporters adding products or export destination.

We contribute to this line of inquiry in two ways. First, we adopt a more flexible estimation strategy than previous studies, and second, we show that the effect of tariff reductions on the extensive margin of trade is particularly strong for differentiated goods. We estimate the tariff elasticity in the context of China's import tariff reductions around the time of its entry into the WTO at the end of 2001. We study Chinese imports rather than exports because many important trading partners already applied Most-Favored Nation (MFN) rates to China's exports already before 2001. Its entry into the WTO made this practice irrevocable and is likely to have had an impact on exports as well, but identifying this effect in a precise way is difficult.[2]

Buono and Lalanne (2012) highlight that the magnitude of the extensive margin response to trade liberalization depends crucially on the nature of the identification strategy. We use product-country rather than firm-level data to avoid relying exclusively on cross-country differences for identification, as in most applications that use the gravity framework. We start from the triple-difference estimator of Frazer and Van Biesebroeck (2010) which identifies the effect of a tariff reduction solely from differential changes over time. They measure the relationship between changes in tariffs and changes in export propensity relative to a baseline probability at the exporter-product level that captures comparative advantage in a disaggregate way. Product-year and country-year dummies control flexibly for variation over time in product-level import demand and country-level export supply. We extend this approach to a quadruple-difference by additionally

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[1] This literature evaluates the impact of both the General Agreement on Tariffs and Trade (GATT) and its successor, the World Trade Organization (WTO). A related literature investigates whether free trade agreements boost their members' trade, see for example Baier and Bergstrand (2007). Haveman et al. (2003) broaden the topic further and include the effects of non-tariff barriers and trade preferences.

[2] A sizable literature exploits the "China shock" (Autor et al., 2016), the fact that from 2001 onward WTO members no longer have discretion which tariffs to apply to China's imports. Any differential impacts across products is assumed to be proportional to the difference in MFN and the average non-MFN rates, even though the benefit to China's exporters of reduced uncertainty is likely to differ across products.

normalizing with exports to a control country at the same level of detail. It guarantees that even changes at the exporter-product level that apply across destinations (importers) are not mistaken for tariff effects.[3] We further incorporate heterogeneity in tariff elasticities by interacting the tariff variable with country or product characteristics.

Our second contribution is to highlight that the tariff elasticity at the extensive product-margin varies systematically across sectors. In particular, it is higher for differentiated goods or goods with a lower import demand elasticity. We provide a theoretical justification for this finding, starting from the generalization of the Melitz-model to many destinations by Chaney (2008). The model predicts the export volume of a new exporter to be negatively related to the elasticity of substitution, but the change in the number of exporters to be independent of the elasticity. However, Besedeš and Cole (2017) point out that this independence does not hold if tariffs are applied not to the marginal cost, but to the final price that includes the markup. While reductions in tariff rates and variable costs have the same effect on a firm's optimal price and quantity, they have a different impact on the absolute level of profit and that is exactly what matters for export market entry.[4]

Our dependent variable for the extensive margin of trade is a dummy variable for the existence of any positive trade flows in a product category. In a heterogeneous firm model with a fixed cost of exporting, this indicator will be one if at least one firm in that product market has a productivity level higher than a threshold that is increasing in the import tariff. No trade flows are observed between many country-pairs, especially if one disaggregates to the product level. This is consistent with the heterogeneous firms model if there is no free entry (Chaney, 2008), if the productivity distributed is bounded (Helpman et al., 2008), or if trade flows are only measured if they exceed a minimum value (Kehoe and Ruhl, 2013). Note that the effect of a tariff reduction on the probability that at least one firm exports necessarily has the same sign as the effect on the number of firms that export. The latter is the usual measure for the extensive margin in studies using firm-level data.[5]

Our first key finding is that the probability of export market entry is highly responsive to tariff declines. This is important for policy as a strong effect at the extensive margin imply that the impact of trade liberalization will be magnified over time as more firms start exporting (Eaton et al., 2008). Our estimates are in contrast with the small effects found for US imports in Debaere and Mostashari (2010) and for French exports in Buono and Lalanne (2012). A possible explanation for the difference could be the much smaller reductions in average tariff rates, respectively 1.4 and 6.4 percentage points over a ten-year period, in those countries. Another

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[3] The estimation approach can also be understood as a triple-difference along the time, importer, and combined exporter-product dimensions. The "within" group change of interest is defined by the interaction of the exporter-product with the importer dimension.

[4] In the model with monopolistic competition and CES demand, prices and marginal costs are proportional. For infra-marginal decisions it does not matter whether tariffs are assessed on the total price or on the marginal cost, e.g. if foreign firms engage in intra-firm trade.

[5] The relative magnitude of the two effects is determined by the shape of the firm-level productivity distribution.

difference is that we work with a more detailed product classification, especially compared to Buono and Lalanne (2012), which makes the extensive margin more prominent.[6] The different identification strategy could be another explanation: we measure changes over time relative to a flexible product-country benchmark while controlling for product and exporter-specific trends and normalizing with export growth to similar destinations.

Secondly, we show that the extensive margin effect is strongest for differentiated goods, using the classification of Rauch (1999) to divide products into three ordered categories: homogeneous, reference-priced and differentiated goods. We assume that this ordering corresponds to a decreasing elasticity of substitution. This ordering for the relative magnitude of effects across products is in line with the heterogeneous firm model of Chaney (2008), if tariffs are assessed on the final good price that includes the markup (Besedeš and Cole, 2017). As a robustness check, we use the Chinese import demand elasticities estimated by Broda, Greenfield, and Weinstein (2006) to show directly that less-elastic goods show a stronger extensive margin response.[7]

To understand the determinants of increased export market penetration, the existing literature has mostly looked at proxies for the fixed cost of exporting. Examples are trade facilitation (Persson, 2013), familiarity with the export destination (Andersson, 2007), or immigration networks (Peri and Requena-Silvente, 2010). In each of these three cases, effects were stronger for differentiated products, consistent with Chaney (2008) who showed that the fixed cost elasticity should decrease with the elasticity of substitution.

A third finding is that the estimated effects are also heterogeneous across country groups. The few countries still outside the WTO in 2001 do not show any response to China's tariff cuts, which supports the effectiveness of our methodology. Moreover, OECD countries show a much larger response than the other WTO members. Perhaps the average demand elasticity for exports from developed economies is on average lower even within each product category. The interpretation of the Melitz-model in terms of quality differentiation by Baldwin and Harrigan (2011) provides a justification for such an assumption. Note that reverse causality, where WTO members negotiate higher tariff reductions for products where they have strong export potential, is unlikely to explain the heterogeneity as the relative elasticities across product categories show the same ordering for non-OECD WTO members which played a minor role at the negotiation table.

The remainder of the paper is organized as follows. Section 2 provides a theoretical framework to help with the interpretation. Section 3 outlines the empirical model and Section 4 the data

6 Buono and Lalanne (2012) find a strong response at the product extensive margin when they use cross-country differences in tariff rates to identify the effects, but not for their "within" estimator that exploits changes over time.

7 The import demand elasticities are used to show robustness, but not for the benchmark results as they are themselves estimated from Chinese import flows and could raise an endogeneity concern. It is a relatively minor issue as import growth at the extensive margin accounts for only a fraction of total import growth and the sample periods in our paper barely overlaps with the period used in Broda et al. (2006). Moreover, to the extent that endogeneity is an issue, it will bias our tariff elasticity estimates towards larger responses in high-elasticity sectors, while our key finding is a negative correlation between the extensive margin effect and the demand elasticity.

description. Estimation results are in section 5, followed by a discussion of potential endogeneity problems and implications in Section 6. Section 7 concludes.

2. Theoretical framework

2.1 Productivity threshold for exporters

The seminal contribution of Melitz (2003) introducing firm heterogeneity in an equilibrium model of international trade features two key elements: firms differ in productivity and they need to pay a fixed cost to enter the export market. In equilibrium, only the sub-set of firms with a productivity level that exceeds a minimum threshold are able to sell profitably abroad. Chaney (2008) generalizes the model to a world with many asymmetric countries and derives predictions for the structure of bilateral trade flows. The adjustment of exports to a change in fixed or variable trade cost is now the combined effect of a change in the volume of trade for existing exporters (the intensive margin) and the exports by firms that enter or leave the export market (the extensive margin). As in the homogenous firms model of Krugman (1980), the adjustment at the intensive margin is increasing in a sector's elasticity of substitution. In contrast, the adjustment at the extensive margin is decreasing in the substitution elasticity only in the heterogenous firms model. The latter effect even dominates if firm-level productivity follows a Pareto distribution.

Chaney (2008) derives the minimum productivity threshold for exporters as:

$$\bar{\varphi}_{cd} = \lambda \left(\frac{Y}{Y_d}\right)^{1/\gamma} \left(\frac{w_c \tau_{cd}}{\theta_d}\right) f_{cd}^{\,1/(\sigma-1)}. \qquad (1)$$

Firms from origin country c in a sector with substitution elasticity σ are predicted to require a productivity level above the $\bar{\varphi}_{cd}$ threshold to at least break even exporting to destination country d. The fixed costs of trading between the two countries is denoted by f_{cd} and the variable costs, which are of the iceberg variety, by τ_{cd}. The remaining variables are the same for all sectors: Y_d / Y is the share of country d in world income, γ is the shape parameter of the Pareto distribution, w_c is the wage in country c, θ_d is the index of remoteness for country d, and λ is a function of several parameters in the model.

Besedeš and Cole (2017) use the same framework but argue that changes in foreign import tariffs (denoted by t_{cd}) affect the level of profits differently from changes in variable trade costs such as transportation costs. In the CES demand model with monopolistic competition, the price-cost markup is constant and a multiplier has the same effect when applied to the marginal cost or to the final price. As a result, a change in t_{cd} or τ_{cd} changes the price and quantity of all exported units in the same way. However, if the tariff is applied to the final price that includes the markup, it will take away more revenue than a variable trade cost that is applied to the marginal cost. In particular, it will also tax away some of the profit markup and it will have a larger impact on extensive margin decision which depends on the absolute profit level. In equilibrium, tariff changes will thus affect the number of exported varieties to a greater extent than variable cost changes. Because the markup is decreasing in the substitution elasticity, the difference is larger for sectors with a low elasticity of substitution. Besedeš and Cole (2017) show that it modifies the productivity threshold for exporters to:

$$\bar{\varphi}_{cd}^{*} = \lambda \left(\frac{Y}{Y_d}\right)^{1/\gamma} \left(\frac{w_c \tau_{cd}}{\theta_d}\right) \left(f_{cd} t_{cd}^{\sigma}\right)^{1/(\sigma-1)}. \tag{2}$$

The comparative statics for any extensive margin response will depend on the effect on the minimum productivity threshold. The elasticities with respect to changes in fixed and variable costs or tariff rates are as follows:

$$\frac{\partial \ln \bar{\varphi}_{cd}^{*}}{\partial \ln f_{cd}} = \frac{1}{(\sigma - 1)} \tag{3}$$

$$\frac{\partial \ln \bar{\varphi}_{cd}^{*}}{\partial \ln \tau_{cd}} = 1 \tag{4}$$

$$\frac{\partial \ln \bar{\varphi}_{cd}^{*}}{\partial \ln t_{cd}} = \frac{\sigma}{(\sigma - 1)} = 1 + \frac{1}{(\sigma - 1)}. \tag{5}$$

We are particularly interested in the dependence of these elasticities on a sector's elasticity of substitution σ (also the demand elasticity). The effect of a change in the fixed cost on the threshold, and hence also on proxies for the extensive margin change, such as the number of exporters, is declining in σ, consistent with the findings in Andersson (2007), Persson (2013), and Peri and Requena-Silvente (2010). In contrast, the response to a change in variable cost does not depend on the substitution elasticity and is predicted be the same for all sectors or product categories. When researchers use a variable like distance as a proxy for trade costs, it is not obvious whether equation (3) or (4) applies. One can argue that distance raises the iceberg transport costs, but overcoming a larger distance has a fixed cost aspect as well.

An ad valorem tariff, on the other hand, is clearly a variable trade cost. The tariff elasticity in equation (5) indicates that the extensive margin response also depends on the substitution elasticity in this case. The tariff elasticity is higher than the other two elasticities, in fact, it equals the sum of the fixed cost and variable cost elasticities (3) and (4). As a result, a tariff reduction will have an additional effect in low-elasticity sectors, like a fixed cost reduction. The intuition is that the absolute profit increase is magnified for low-elasticity goods as the original tariff is also applied to the profit markup. As a result of the higher profits, more firms will be tempted into the export market, especially in differentiated goods sectors which are assumed to be of low elasticity. The intensive margin effect of a tariff reduction, lowering consumer prices and raising sales, is the same as that of a variable cost reduction and increases with the elasticity of substitution.

2.2 The extensive margin

Chaney (2008) defines the extensive margin response as the volume of trade by new exporters. It is the product of the measure of new firms entering the export market and the total export sales that they realize. The decrease in the extensive margin effect with a product's substitution elasticity that he highlights is not driven by the number of new exporters, but by their larger export volume in low-elasticity sectors. New exporters are by construction the least productive firms in the export market. When product differentiation is low and the demand elasticity is high, they will have great difficulties competing with more productive competitors and only sell small volumes. It is exactly the stronger intensive margin response of incumbents in high-elasticity sectors that lowers the export sales of new exporters which is included in his definition of the extensive margin effect.

Eaton et al. (2008) empirically quantifies the importance of the extensive margin defined exactly as in Chaney (2008). Kehoe and Ruhl (2013) adopt a similar definition to study the importance of previously "untraded" product categories in several trade liberalization episodes. They use a liberal definition of untraded products, combining all 5-digit product codes in the SITC classification that together account for less than 10% of the total trade volume of a country-pair. They show that trade volumes for this group of products, which combines a pure extensive margin response with an intensive margin response for new or marginal exporters, increases disproportionately.

They adopt that particular definition because reporting thresholds for trade transactions vary across countries and are bounded away from zero. In some cases, they even vary over time or by product category. To deal with this variation, Evenett and Venables (2002) set a higher, but uniform threshold and only classify exporter-importer-product relationships as non-zero if a trade flow exceeds a value of $50,000 in a given year. Most of the literature, however, uses the trade data as reported, classifying any reported trade flow as a positive trade relationship and all reported zeros as no-trade relationships.

Researchers working with firm-level data can relate directly to the theory and simply measure the change in the number of new exporters following a change in tariffs or variable costs, see for example Buono and Lalanne (2012). Researchers working with product-level data for country pairs are forced to take a different approach. Often, they first show theoretical results in the standard heterogeneous firms framework assuming single-product firms in a home and one or more foreign countries. When they turn to estimation, they make the Armington assumption and equate unique product varieties (firms) in the model with imports from different countries. They measure the change in the number of trading partners for a particular product as a proxy for the extensive margin effects, see for example Arkolakis et al. (2008) and Helpman et al. (2008). Of course, if one importer is replaced by another, this measure will not record any extensive margin change.

With free entry and an unbounded productivity distribution, there will be a positive probability of exporting in each product category for any country pair. Using such a literal interpretation of the model, no changes in the number of trading partners should ever be observed. That is clearly at odds with the data. Helpman et al. (2008) show that even at the aggregate level around half of all country-pairs do not show any positive trade flows. One does not have to look at a very disaggregate product level for the trade matrix to be entirely dominated by zeros. The literature has used several approaches to make the heterogeneous firms model consistent with the observed preponderance of zero trade relationships. Chaney (2008) eliminates the free entry assumption, taking the number of firms, and thus potential exporters, as given. Helpman et al. (2008) assumes that the productivity distribution is bounded from above. Melitz and Ottaviano (2008) use a quasi-linear demand system that, in contrast with CES preferences, has bounded marginal utility when consumption of a variety approaches zero.

We also work with product-level bilateral trade data, but to stick closer to the theory, we use a different dependent variable. Our empirical specification models the probability that a positive export flow is observed. This will be the case when the most productive firm in a sector has entered a particular export market. If a country pair had no trade flows for a particular product prior to the tariff reduction, but a positive flow is observed afterwards, we know that at least one active firm must now have a productivity level above the required threshold. This definition ignores how much each new exporter sells, and is thus a pure extensive margin effect.

The comparative static for how this measure—in expectation—varies with the import tariff is given by:

$$\frac{\partial \ln Prob(x \geq \bar{\varphi}_{cd}^{*})}{\partial \ln \bar{\varphi}_{cd}^{*}} \frac{\partial \ln \bar{\varphi}_{cd}^{*}}{\partial \ln t_{cd}} = -\gamma \left(1 + \frac{1}{\sigma - 1}\right). \tag{6}$$

Assuming firm-level productivity follows the Pareto distribution, the first term in the derivative simplifies to $-\gamma$, with γ the shape parameter.[8] As a result, the change in probability that any firm has a productivity level that exceeds the minimum threshold $\bar{\varphi}_{cd}^{*}$ and enters the export market varies only across sectors through the comparative statics discussed previously. In particular, the tariff elasticity declines (in absolute value) with the elasticity of substitution σ.

3. Empirical model

To identify the tariff elasticity of exports as cleanly as possible, we augment the triple difference-in-differences approach of Frazer and Van Biesebroeck (2010) where we have replaced the product dimension by an exporter-product pair and use destinations as the third dimension. The dependent variable is an indicator which equals one if we observe positive exports of product *p* from country *c* to country *d* in year *t*, and equals zero otherwise. The key explanatory variable is the (logarithm of the) import tariff that applies to this trade flow and its coefficient captures the extensive margin response to a tariff decline.

[8] The cumulative density function for the Pareto distribution is $Prob(x \leq \bar{x}) = 1 - \left(\frac{x_m}{\bar{x}}\right)^{\gamma}$. The result in equation (6) does not depend on the scale parameter x_m which is usually normalized to 1.

We estimate the equation with least squares, which implies a linear probability model, as it allows the inclusion of extremely flexible controls for demand and supply effects.[9] The expectation, of the dependent variable is the probability that the productivity of the most efficient firm exceeds the threshold $\bar{\varphi}^*_{cd}$. The elasticity of substitution differs by product and the productivity dispersion and the wage rate differ across product-exporter combinations. In addition, fixed and variable trade costs are likely to be product and country-pair specific and vary over time. To control flexibly for all these factors, we include a large number of three-way interaction fixed effects exploiting the four dimensions of the data, i.e. exporter, destination, product and time.[10]

For the baseline results, we estimate a separate regression for each product category *k* of the Rauch classification:

$$I\left[Export_{cdpt} > 0\right] = \beta_k \ln\left(1 + t_{cdpt}\right) + \gamma_{cdp} + \gamma_{cpt} + \gamma_{cdt} + \varepsilon_{cdpt}\ . \qquad (7)$$

A unit of observation is a particular product *p,* year *t* and country pair *cd,* measuring exports from 124 exporter countries *c* to each of 3 destination countries *d*. China is one of the three destination countries and we use India and Indonesia as control destinations. Almost every product imported in China has experienced a tariff reduction over the 2000-2006 period, while import tariffs of India and Indonesia do not vary over the sample period. The tariff elasticity is then identified by the strength of the association between Chinese product-level tariff cuts and increased import penetration in China relative to changes in the control destinations.

The fixed effects included in the regression are crucial. Exporter-product-year effects soak up product-specific changes in demand, quality, and production technology that are global in nature. We have chosen India and Indonesia as control destinations as we expect the evolution of their import demand to be somewhat similar to China's, because they are also WTO members, important manufacturing bases in Asia, large developing countries, and located at a similar distance from important exporters. Country-pair/year effects soak up changes in the national business cycle, level of development, trade costs, and overall trade expansion due to greater integration in the world economy. Country-pair/product effects capture the baseline probability of bilateral trade for every product, including the effect of Indian or Indonesian tariff rates. We do not aim to explain the *level* of import penetration, only to what extent relative *changes* over time can be explained by Chinese tariff reductions.

The parameter of interest on the logarithm of import tariffs measures the elasticity of the probability that a product is exported to China with respect to the import tariff. In this specification, we implicitly assume a constant tariff elasticity, but we also estimate equation (7) using the level of tariffs instead, which assumes a constant effect of each percentage point reduction, i.e. a

[9] As a robustness check we also estimate a probit model that incorporates the discrete nature of the dependent variable, but does not allow as flexible controls.

[10] Only the destination-product-year fixed effect cannot be included as it would absorb all import tariff effects.

constant semi-elasticity. k indexes the Rauch classification's product categories which are expected to differ in their elasticity of substitution. The different β_k parameters are estimated by interacting the tariff variable with category dummies. According to the model in Besedeš and Cole (2017), we expect the tariff elasticity to be the larger in absolute value if the elasticity of substitution for the product category is lower.

The intuition for equation (7) can be illustrated by explicitly writing out the triple-difference, over each dimension: time, exporter-product, and destination:

$$\left(\Delta IM_{C,1} - \Delta IM_{C,2}\right) - \left(\Delta IM_{I,1} - \Delta IM_{I,2}\right) = \beta\left(\Delta \ln(1+t)_{C,1} - \Delta \ln(1+t)_{C,2}\right) \quad (8)$$

For two exporter-products (1 and 2) and two export destinations (C and I) we calculate the change over time in the probability that the exporter-product-destination observation shows positive exports. The relative increase in import penetration for product 1 versus 2 in China is normalized by the relative increase in import penetration in India or Indonesia. Given that treatment here is continuous and not a zero-one decision, we normalize the double-difference in growth rates by the relative change in Chinese import tariffs for product 1 versus 2 to obtain the elasticity estimate β.[11] The corresponding difference for India or Indonesia in the right-hand side drops out as their tariffs were constant.

The advantage over a standard difference-in-differences (DID) estimation at the country or product level is the robustness to policy endogeneity (Besley and Case, 2000) and to misattribution of effects from omitted variables to the policy variable. The advantage over a country-level DID estimation is that anything unusual about the evolution of China's trade will be differenced out in the product dimension. The advantage over a product-level DID estimation for a single destination is that strong performance of product 1 exports to China for other reasons than tariff reductions are normalized by the relative export performance to India and Indonesia. Debaere and Mostashari (2010) and Buono and Lalanne (2012) also use data sets with observations defined as country-product-time triplets, but they only include additive fixed effects. As a result, much of their identifying power still comes from variation in absolute growth rates of import penetration at the country-product level. Buono and Lalanne (2012) illustrate that in their case, the contribution of the extensive margin effect disappears if they use a more flexible triple-difference set-up.

Following China's entry in the WTO, its product-level import tariffs declined to different extents. We allow the export responses to differ across three categories of homogeneous, reference-priced, and differentiated goods, as defined by Rauch (1999). For the baseline results, we estimate equation (7) on observations for year 2001 and 2006, we include both India and Indonesia as control destinations, and use the conservative classification of Rauch (1999) to define product categories. In robustness checks, we vary the sample period, the control destinations, use Rauch (1999)'s liberal classification, or import demand elasticities estimated by Broda et al. (2006)

[11] The right-hand side in equation (8) approximates the percentage difference $(\Delta t/t_0)_{C,1} - (\Delta t/t_0)_{C,2}$.

to categorize goods; we also estimate a probit model. We first pool across all exporters, assuming constant tariff elasticities across countries. Subsequently, we relax this and estimate the model separately for OECD countries, countries that are members of the WTO but not the OECD, and countries belonging to neither group.

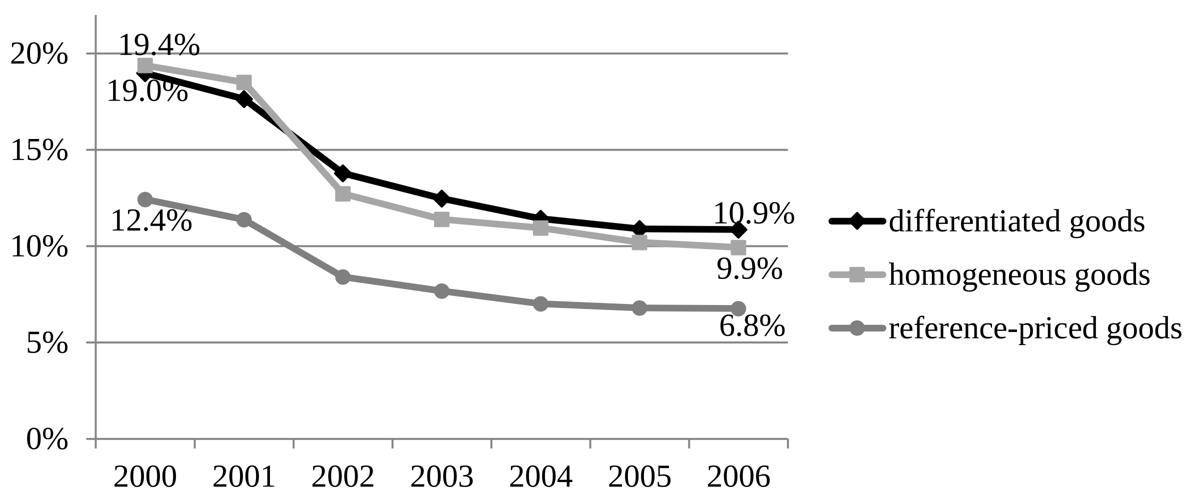
4. Data

4.1 Chinese import tariffs

We obtained Chinese import tariffs at the 8-digit level of the Harmonized System (HS) classification of goods from the Chinese National Bureau of Statistics. Because we use bilateral trade flows between different country-pairs, e.g. US-China and Germany-India, we need to work at the 6-digit HS level, the most detailed product categories that are standardized across countries.

Ideally, tariff rates should be aggregated using weights that are unimpeded by protectionist barriers. Given that many 6-digit categories contain only a single 8-digit sub-category and when there are several sub-categories their tariff rates tend to be rather similar, the exact choice of weight is not very important, but we want to avoid introducing a systematic bias. In particular, using current import flows as weights would bias the average tariff down, especially in early years. One possibility is to construct weights based on information from the last year in the sample—when tariff rates are lowest—and apply a constant weight over the entire period. A problem is that the 8-digit HS classification underwent several changes between the 1996 and 2002 versions. Therefore, this would risk introducing tariff changes that are merely due to reclassifications. Because our within-group identification strategy is particularly susceptible to spurious tariff changes, we instead opted for a simple, unweighted average.

Figure 1: Evolution of average Chinese import tariffs by products category



Notes: Unweighted average of Chinese import tariffs at the HS 6-digit level. Products are classified using Rauch's (1999) conservative standard.

The evolution of the average Chinese import tariffs from 2000 to 2006 is depicted in Figure 1. The average tariff declines significantly for all three product categories. Chinese tariffs had already started to decline before 2000, but we only rely on tariff cuts from China's WTO accession at the end of 2001 onwards. Between 1995 and 2000 tariffs also fell in other countries, as the outcome of the Uruguay Round agreement. However, over the 2000-2006 period, import tariffs for India and Indonesia remained almost perfectly constant. All identifying power comes from product-level variation in the Chinese tariff reductions.

4.2 Bilateral trade flows

The dependent variable is an indicator for positive imports in one of the destination countries between 2000 and 2006 for products defined at the 6-digit HS level. Indian and Indonesian trade data are taken from the BACI database of CEPII.[12] Chinese trade data are from China Customs and is aggregated up from the firm-level. It allows us to exclude imports that enter the country under the export processing trade regime and are not subject to import tariffs. Exporting countries are included in the sample if they show any positive exports to at least one of the three destinations in the first or last year of the sample. The list of 135 countries is provided in Table A.1 in the Appendix. They are classified in three groups depending on membership in the WTO and/or OECD.

We limit the sample to manufactured products, omitting agricultural and mining products. To classify them in categories that differ in elasticity of substitution, we adopt the three-way classification of Rauch (1999) which is at the 4-digit SITC level.[13] The final sample contains 3,378 manufactured goods across the three categories, out of a total of 5,040 (6-digit) products in the BACI data set. Rauch (1999) developed a conservative and liberal classification and the breakdown of products for both cases is shown in Table 1. Most products fall in the category of differentiated goods, while homogenous products are relatively rare, representing only 4 to 5.7 percent of manufactured goods in our sample.

As a robustness check, we capture product heterogeneity using the estimated import demand elasticities for China taken from Broda et al. (2006). They are provided at the 3-digit HS level (for a total of 170 categories) and have a median elasticity of 3.4, mean of 6.2, and standard deviation of 11.9. We use them to classify all 6-digit products into three equally sized groups of low, medium, and high elasticity products. We also use the elasticity estimates directly, interacting them with the tariff rates.

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[12] BACI reconciles the declarations of exporters and importers in the United Nations' COMTRADE database. It extend considerably the number of country-pairs for which trade data are available. Details on the harmonization procedure are provided in Gaulier and Zignago (2010) and the data set itself is available online at http://www.cepii.fr/anglaisgraph/bdd/baci.html.

[13] A concordance table between the HS and SITC classifications is used to link this information to the trade data. Both the Rauch classification and concordance table are available at:
http://www.macalester.edu/research/economics/page/haveman/Trade.Resources/Tradeconcordances.html

Table 1: Breakdown of the sample by product category

	Conservative classification		Liberal classification	
	Number	Fraction	Number	Fraction
Differentiated goods	2,336	69.1%	2,133	64.8%
Referenced-priced goods	908	26.9%	972	29.5%
Homogeneous goods	134	4.0%	185	5.7%
Total	3,378		3,290	

With the set of countries, products, and sample years determined, the panel is balanced along all three dimensions by adding zeros for missing trade flows. The number of observations in each regression is then the product of the number of years (2001 and 2006), by the number of exporting countries (124 WTO members), import destinations (China, India, and Indonesia), and products (3,378 6-digit HS categories). In the baseline analysis, we use 2,513,232 observations: 1,737,984 differentiated goods, 675,552 reference-priced goods, and 99,696 homogeneous goods.

Table 2 lists the number of products for which we observe positive imports by year, product category, and export destination. The initial import penetration varies between 7.0 and 8.4 percent of all potential exporter-product pairs, leaving a lot of scope for growth at the extensive margin of trade. In the following six year, the import penetration increases everywhere, but most strongly for China and for differentiated goods.

Table 2: Number of positive import flows

	China			India			Indonesia		
	2000	2006	Change	2000	2006	Change	2000	2006	Change
Differentiated	26,068	38,309	+47.0%	28,711	39,843	+38.8%	23,506	27,025	+15.0%
Reference-priced	7,265	9,227	+27.0%	8,905	10,853	+21.9%	7,625	7,620	–0.1%
Homogeneous	549	831	+51.4%	722	952	+31.9%	642	637	–0.8%
Total	33,882	48,367	+42.8%	38,338	51,648	+34.7%	31,773	35,282	+11.0%
(% of potential)	(7.4%)	(10.6%)		(8.4%)	(11.3%)		(7.0%)	(7.7%)	

Notes: The counts are the number of product-country pairs with positive imports into the respective countries. Percentages at the bottom are expressed as a fraction of the potential number of trade flows (3378 products x 135 exporters). Products are classified using the conservative classification of Rauch (1999).

5. Empirical Results

Baseline estimates of equation (7) are reported in Table 3. This uses data for 2001 (the year China entered the WTO) and 2006, 124 exporters (all WTO members), 3 destinations, and all manufactured products that could be matched to the classification of Rauch (1999). Results in the first and third columns are tariff elasticities, estimated in the form suggested by the theory, interacting the three product category dummies with ln(1+*t*). Results in the second and fourth columns are semi-elasticities, where tariff rates enter linearly. The patterns are the same in both cases.

In the model of Chaney (2008), the tariff elasticity at the intensive margin ($\partial \ln EXP / \partial \ln(1+t)$) equals $(\sigma - 1)$. The estimates in column (1) imply that the elasticity of substitution is higher for more homogeneous products and equals 1.5, 2.8, and 7 for the three product categories. Both the ordering and the magnitudes are very reasonable. The literature estimating the trade elasticity—the responsiveness of trade flows to changes in marginal costs or prices—is vast. An important finding is that point estimates vary greatly when aspects of the specification or identification are changed (Boehm et al. 2020). For example, estimates are much larger exploiting long-run versus short-run variation, which helps explain the higher numbers obtained in the trade literature than in the macro literature. Moreover, the absolute magnitudes tend to be lower if more flexible controls are included.

Table 3: Benchmark estimates of tariff elasticities at the extensive and intensive margins

Dependent variable:	(log) Export value (intensive margin)		Dummy for positive exports (extensive margin)	
	(1)	(2)	(3)	(4)
ln(1+*tariff*)				
* Differentiated product	-0.488		-0.092***	
	(0.381)		(0.010)	
* Reference-priced product	-1.775***		-0.037***	
	(0.613)		(0.010)	
* Homogenous product	-6.082***		-0.026**	
	(1.958)		(0.012)	
tariff				
* Differentiated product		-0.498		-0.079***
		(0.324)		(0.006)
* Reference-priced product		-1.404***		-0.029***
		(0.445)		(0.008)
* Homogenous product		-4.166***		-0.017**
		(1.415)		(0.007)
Observations	110,398	110,398	2,513,232	2,513,232

Notes: Sample period is 2001 & 2006; export destinations are China, India, & Indonesia; products are classified using the conservative classification of Rauch (1999). ***, **, and * indicate significance at the 1%, 5% and 10% level.

For our purpose, the most important pattern of the estimates in columns (1) and (2) are the much higher estimates (in absolute value) for homogenous compared to differentiated goods. This pattern is in line with results in Tang (2006), but the absolute magnitudes are not comparable as she includes linear and quadratic absolute tariff levels as explanatory variables. While the estimates are consistent with the theoretical prediction, the benchmark model naturally omits many factors. For instance, Javorcik and Narciso (2008) also find trade in differentiated goods to be less responsive to tariff changes, but they explain this by greater scope for tariff evasion.

Results for the responsiveness at the extensive margin in columns (3) and (4) are the primary interest. All point estimates are negative and significant at the 1 or 5 percent level. This is in contrast with the existing literature that tends to find only small and mostly insignificant effects of tariff reductions on the extensive margin. Most importantly, the elasticity is much higher in absolute value for differentiated goods which are assumed to have the lowest demand elasticity, followed by reference-priced goods, and then the homogeneous goods. This pattern is the inverse of the ranking at the intensive margin and is consistent with the theoretical prediction. Note that the dependent variable measures the first export instance within a 6-digit product category. The existence of several varieties within a particular category does not invalidate our empirical strategy, but it might bias down the effects.

Results in Table 4 illustrate that the estimates at the extensive margin are robust to changing several elements of the benchmark specification. In columns (1) and (2), each of the control destinations is used separately. In column (3), the control destinations are omitted and we estimate a triple-difference specification. In column (4), we use the liberal classification of Rauch (1999) which classifies fewer products as differentiated. In column (5), we start the sample period already in 2000, two years before the negotiated tariff cuts start taking effect. Tariff cuts in the pre-WTO period were voluntarily and they could be concentrated in sectors where the scope for increased import penetration was low.[14] In column (6) we include observations in all intervening years. Finally, in column (7), we report coefficient estimates for a probit model. This makes it impossible to include the rich set of fixed effects, but we include a random effect at the product level, as in Baldwin and Harrigan (2011).

A few patterns are worth highlighting. Most importantly, the qualitative pattern is highly robust. The estimate for differentiated products is always the largest and highly significant in all but one instance. Results for the other two product categories are always lower in absolute value and sometime become insignificant. In six of the seven columns, the elasticity for reference-priced goods, the intermediate category, is more negative than for homogenous goods. The only case where the point estimate for differentiated goods becomes insignificant is when all intervening years (all 6 years from 2001-2006) are included. The inclusion of many instances where tariffs do

[14] Data on ordinary trade separately from duty-free, processing trade is only available from 2000 onwards.

Table 4: Robustness checks

	Dependent variable is a dummy for positive exports						
	Only India as control destination	Indonesia as control destination	No control destinations (triple-diff.)	Liberal product classification	Extended period: 2000-2006	Include all intervening years	Probit estimation
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ln(1+*tariff*)							
* Differentiated	-0.081***	-0.103***	-0.206***	-0.088***	-0.110***	-0.026	-0.062***
	(0.008)	(0.010)	(0.061)	(0.009)	(0.007)	(0.041)	(0.007)
* Reference -priced	-0.041***	-0.034**	0.006	-0.052***	-0.024***	-0.008	-0.015
	(0.010)	(0.014)	(0.038)	(0.011)	(0.009)	(0.018)	(0.010)
* Homogeneous	-0.040***	-0.012	-0.024	-0.035**	-0.019	0.022	-0.0003
	(0.010)	(0.012)	(0.018)	(0.016)	(0.015)	(0.026)	(0.012)
Observations	1,675,488	1,675,488	959,352	2,447,760	2,513,232	7,340,304	1,256,616

Notes: In the benchmark estimation, the sample period is 2001 and 2006, export destinations are China, India, and Indonesia, and products are classified using the conservative classification of Rauch (1999). In each column one of these dimensions is changed. ***, ** and, * indicate significance at the 1%, 5% and 10% level.

not change lead to much higher standard errors in our within estimator.[15] The results in column (3) confirm the findings in Buono and Lalanne (2012) and Boehm et al. (2020) that the estimated responses are a lot higher when cross-sectional variation contributes to the identification. Hence, our baseline results can be considered conservative.

In Table 5, we allow the elasticities to differ not only by product category, but also by country grouping. The ordering of the elasticities, highest for differentiated and lowest for homogenous goods, applies to OECD countries as well as non-OECD countries. Within each product category, exports from OECD countries are always more elastic. The difference is especially large for differentiated products: a point estimate of –0.221 versus –0.045. One explanation is that exporters from OECD countries face a lower average substitution elasticity even within each category, for example because they sell higher quality goods (Schott, 2004). According to this interpretation, it is natural that the estimates differ less for homogenous goods: the point estimates are –0.043 versus –0.020 and are not significantly different.[16]

[15] The same is true comparing the results in Rose (2004) and Subramanian and Wei (2007): including intervening years notably lowers the significance of the estimates. We also obtained results that confirm a delayed effect of WTO entry, as in Chang and Lee (2011), but the need to observe ordinary and processing trade separately again prevents us from extending the sample period beyond 2006.

[16] Another distinction is that firms in non-OECD countries had to confront fiercer competition in their home markets as Chinese products could now be imported at lower MFN tariffs. With heightened competition at home, expanding abroad might not be the first concern for these firms. In contrast, the United States and the EU lowered tariffs on Chinese imports long before China’s entry in WTO.

Table 5: Results by country group

	Dependent variable is a dummy for positive exports		
	OECD countries (WTO members)	Non-OECD countries (WTO members)	Non-WTO members
	(1)	(2)	(3)
ln(1+*tariff*)			
* Differentiated product	-0.221***	-0.045***	-0.007
	(0.050)	(0.008)	(0.015)
* Reference-priced product	-0.063**	-0.028***	0.000
	(0.025)	(0.005)	(0.000)
* Homogeneous product	-0.043**	-0.020*	0.015
	(0.020)	(0.011)	(0.023)
Observations	668,844	1,844,388	222,948

Notes: Same specification as as in the benchmark case: sample period is 2001 and 2006, export destinations are China, India, and Indonesia, and products are classified using the conservative classification of Rauch (1999). ***, **, and * indicate significance at the 1%, 5% and 10% level.

In the last column of Table 5 we report results for exports by non-WTO members. The tariffs that China applies to these countries did not have to follow the changes in its MFN tariff rates. This can be considered a placebo test for our identification strategy. It is reassuring, but not surprising that China's WTO entry had no effect on their extensive margins. All three estimates are now statistically insignificant, indicating that firms from these countries did not increase their import penetration following China's entry into the WTO. The estimates in Table 5 confirm that the WTO promotes trade strongly, but unevenly across countries (Subramanian and Wei, 2007).

For the next set of results, we rely on estimates of the import demand elasticities for China by Broda et al. (2006) as an alternative to the classification of Rauch (1999) to measure product heterogeneity. The results in Table 6 include coefficient estimates for the interaction between the elasticity estimates and the tariff variable. In all three cases in panel (a), for the full sample, the coefficient estimate on the interaction term indicates that the absolute value of the extensive margin response is decreasing in the demand elasticity. This is again consistent with the theoretical prediction that tariff elasticities should be decreasing in the elasticity of substitution between products.

Results in panel (b) are for the same specification, but now estimated separately for each category of differentiated products. The difference in estimated tariff elasticity for products with a lower demand elasticity is far more pronounced on this sample. The positive coefficients on the interaction term are approximately twice as high as on the full sample and they are always statistically significant at the 1 percent level.

Table 6: Results with import demand elasticity interaction

	Dependent variable is a dummy for positive exports		
	All countries (1)	OECD countries (2)	Non-OECD countries (3)
(a) All products			
ln(1+*tariff*)	-0.078***	-0.155***	-0.050***
	(0.004)	(0.024)	(0.007)
ln(1+*tariff*) * demand elasticity	0.0014***	0.0018*	0.0013***
	(0.0005)	(0.0011)	(0.0003)
Observations	2,473,800	658,350	1,815,450
(b) Only differentiated products			
ln(1+*tariff*)	-0.113***	-0.254***	-0.062***
	(0.010)	(0.051)	(0.009)
ln(1+*tariff*) * demand elasticity	0.0026***	0.0039***	0.0022***
	(0.0005)	(0.015)	(0.0002)
Observations	1,698,552	452,034	1,246,518

Notes: The tariff variable is included uninteracted and interacted with the import demand elasticities for China as estimated by Broda et al. (2006). Same specification as in the benchmark case: sample period is 2001 and 2006, export destinations are China, India, and Indonesia. In panel (b), differentiated products are identified using the conservative classification of Rauch (1999). ***, **, and * indicate significance at the 1%, 5% and 10% level.

The relative tariff elasticity for OECD and non-OECD countries depends on both coefficients. While the extensive margin response falls more with the demand elasticity for OECD countries, no reasonable value (for the demand elasticity) can overturn the large difference on the uninteracted tariff coefficient. On both samples, the tariff elasticity is estimated much larger for OECD countries.

We can also use the import demand elasticities to divide products evenly into low, medium, and high elasticity groups and estimate a different response for each group. These results, reported in Table A.2 in the Appendix, confirm that the response is strongest for low-elasticity products. The point estimates for that category are –0.123, –0.243 and –0.079 for the full sample, OECD and non-OECD countries, respectively. These results are similar to those for differentiated products in Tables 3 and 5, which were –0.092, –0.221 and –0.045 respectively. On each sample, the ordering of the three estimates follows the inverse ordering of the demand elasticity.[17]

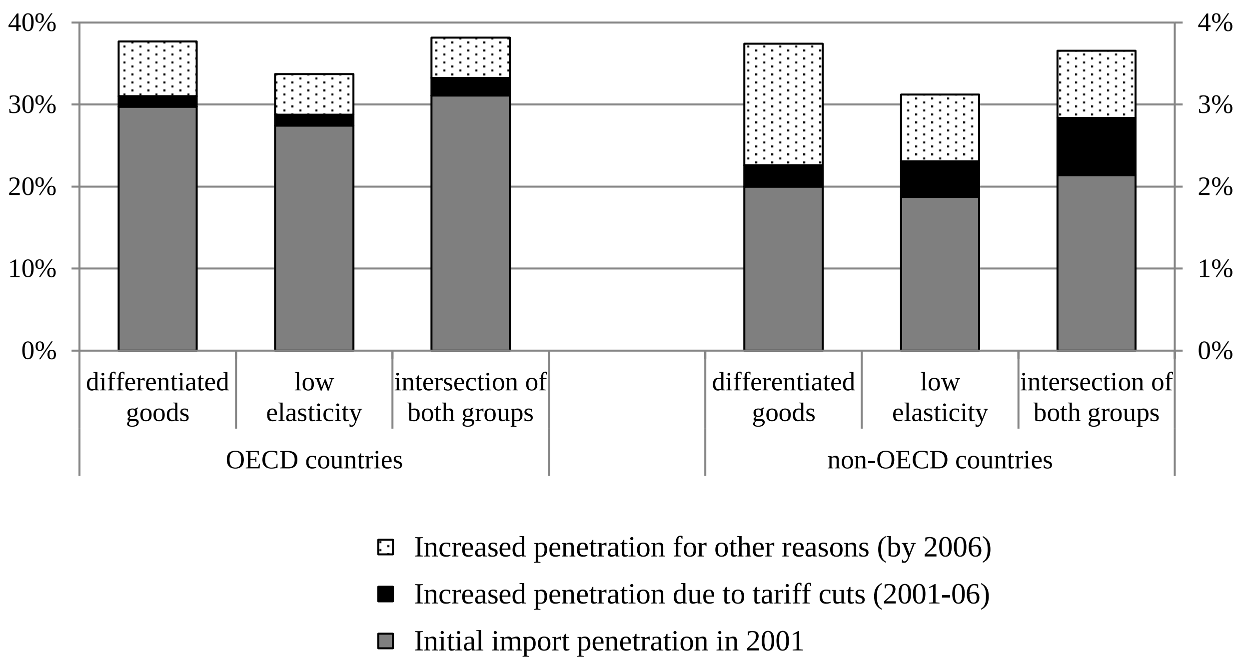
Figure 2 illustrates the economic magnitude of the effects by applying the estimated tariff elasticities to the initial import penetration. The average effects can be calculated directly from

[17] Only for OECD countries, the point estimate is higher in absolute value for the high elasticity than for medium elasticity group, but the difference is not statistically significant.

equation (8) based on the point estimates, tariff cuts and initial penetration rates. Because the average import penetration varies greatly between OECD and non-OECD countries, we report their results side-by-side with a different scale on the two vertical axes. To provide further context, we also show the total change in import penetration between 2001 and 2006. That way, one can see directly what fraction of the total increase in import penetration the tariff cuts account for.

The impact of a tariff cut is largest for the subset of differentiated products with a low demand elasticity. The average decline in import tariffs is 0.062 (in log-points) and the estimated effect of –0.352 for OECD countries implies an increase in the probability of exporting to China of 2.2 percentage points. This accounts for almost one third of the total increase in import penetration of 7.1 percent (up from 31.1 percent in 2001 to 38.2 in 2006). The estimated effect is a lot lower for non-OECD countries, at –0.113, but so is the initial import penetration. The tariff cuts are now estimated to have raised import penetration by 0.7 percentage points, which accounts for almost half of the total increase. The effects are more modest when considering only differentiated goods, with the tariff cuts accounting for approximately one sixth of the increased import penetration, respectively 16.3 and 15.2 percent for OECD and non-OECD countries. Nevertheless, these fractions are substantially higher than the 5 and 12 percent obtained by Debaere and Mostashari (2010) over the 1989-1999 and 1996-2006 periods.

Figure 2: Estimated impact of tariff reductions on average import penetration



Notes: Mean values of the import penetration effects, calculated based on estimates reported in Table 3 and Table A.2. Differentiated products are defined according to the conservative definition of Rauch (1999); low elasticity products are below the median of the Chinese import demand estimates of Broda et al. (2006). The third category are products included in both of these two product categories.

6. Discussion

The findings are consistent with the heterogeneous firm model with horizontally differentiated goods of Chaney (2008) and Besedeš and Cole (2017). The higher tariff elasticity at the extensive margin for goods that face a less elastic demand is driven by their higher profit markup. Following a tariff decline, the equilibrium price adjusts similarly everywhere, but it raises the absolute amount of profit more for low-elasticity, high-markup goods. This makes it easier to cover the fixed cost of exporting and entices more firms into the export market. We now explore alternative explanations due to factors that are omitted from the model.

A first concern is that market entry might be more difficult for homogenous goods, simply because many countries do not produce them. The benchmark results are for a balanced panel, which implicitly assumes that all countries can potentially export each of the 3,378 manufacturing products to China. Agricultural and mining products were already omitted as they are certainly not produced everywhere due to comparative advantage. Markets for products with a high elasticity of substitution are likely to be highly competitive, with low markups and few surviving producers. We expect fewer active firms and potential entrants in such sectors. In theory, the productivity distribution has a positive density over its entire domain in each country and sector, but with indivisibilities there will be country-sector pairs with no active firms, not even in the domestic market.

Table 7: Results using only confirmed exported products for each country

	Dependent variable is a dummy for positive exports		
	All countries (1)	OECD countries (2)	Non-OECD countries (3)
ln(1+*tariff*)			
* Differentiated products	-0.113***	-0.221***	-0.059***
	(0.016)	(0.065)	(0.019)
* Reference-priced products	-0.051**	-0.062*	-0.044**
	(0.023)	(0.036)	(0.018)
* Homogeneous products	-0.040	-0.029	-0.037
	(0.029)	(0.035)	(0.030)
Observations	880,740	324,891	555,849

Notes: Sample is an unbalanced panel in the product dimension. For each exporter country, it only includes products that are ever observed as exports (to any country) over the sample period. Same specification as in the benchmark case: sample period is 2001 and 2006, export destinations are China, India, and Indonesia, and products are classified using the conservative classification of Rauch (1999). ***, **, and * indicate significance at the 1%, 5% and 10% level.

To take into account that some countries may never export certain products, we drop from the sample all country-product pairs with no exports to any destination over the 2000-2006 period.[18] Results in Table 7 illustrate that the basic pattern survives in the reduced sample; the estimates for OECD countries are almost unchanged. For non-OECD countries, the estimates become more negative, as expected, but the change is similar for all three product categories. For homogenous goods the point estimate goes from –0.020 to –0.037, and for heterogeneous goods from –0.045 to –0.059. The lack of exports for some goods is primarily an issue for poorer countries, but it is not concentrated among homogenous products.

A second concern is the potential endogeneity of tariff reductions which were decided in negotiations between China and the existing WTO members. It is possible that countries pushed harder for reductions in product categories where they expected a strong export potential, for example because of comparative advantage (Subramanian and Wei, 2007). This is especially a concern for the OECD countries, as the United States and the EU were the most prominent negotiation partners and other OECD countries are likely to have similar comparative advantage. Such endogeneity could contribute to the higher estimates for differentiated goods as OECD exports are concentrated in that category (Brandt et al., 2017).

There are several reasons why this factor is unlikely to explain much of the effects. Evidence in Blonigen and Cole (2011) demonstrates that since 1993 the presence of FDI in a sector is a better predictor of Chinese import tariffs than foreign export elasticities. Brandt et al. (2017) further illustrate that tariff cuts between 1995 and 2006 are almost perfectly proportional to the initial tariff rate with very little discretion in rate reductions. The averages in Figure 1 also indicate similar tariff declines in each of the three categories. Given that non-OECD countries were less involved in the negotiations and that their comparative advantage is likely to differ from that of OECD countries, Chinese tariff reductions can be considered relatively exogenous from their perspective. It is reassuring that the estimates for the three product categories follow the same pattern also for these countries.

The third element absent from the model is quality differentiation. To the extent that higher quality goods simply enjoy a lower elasticity of substitution, the mechanism we focus on encompasses quality effects. However, there might be a different mechanism too. Baldwin and Harrigan (2011) illustrate that in the Melitz model quality differences can lead to differentiation even within a sector. They predict that, all else equal, the highest quality goods will be shipped the farthest. Hummels and Skiba (2004) derive a similar prediction and provide empirical evidence, starting from the existence of per-unit shipping costs, i.e. not purely of the iceberg variety.

Given that tariffs, like distance, raise the variable trading costs, countries with higher tariffs will disproportionately receive imports of higher quality. Lowering tariffs just enough to entice

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[18] Estimation on the unbalanced panel is a lot computationally intensive, because we now need to perform the within-transformation also on the interaction fixed effects (see Van Biesebroeck and Frazer, 2010). The Stata command “reghdfe” instead implements an iterative procedure to estimate the model.

the first firm to start exporting is similar (hypothetically) to lowering the distance just enough to make exports profitable. For different reasons, both Hummels and Skiba (2004) and Baldwin and Harrigan (2011) predict that the first product exported from exporter *c* to destination *d* will be a high quality product. It is consistent with our finding of a stronger response for OECD exports, which tend to produce more high quality goods (Schott, 2004). Still, the fact that that the ordering of tariff elasticities across product categories also holds for non-OECD countries, which are unlikely to be the quality leaders in all but a few examples, suggests again that quality is unlikely to be the entire explanation.

A final element missing from the model are variable markups. Marginal exporters have lower productivity and higher marginal cost than infra-marginal exporters. With a more flexible demand than the CES, marginal exporters are predicted to choose a lower markup as they face a more elastic residual demand. The difference between marginal and infra-marginal firms is especially pronounced in industries with low elasticity of substitution. It is an open question whether a differential impact on markups reinforces or diminishes the dependency of the extensive margin response on the demand elasticity, as derived for the CES demand case. The analysis in Melitz and Ottaviano (2008) highlights the important role of general equilibrium effects, making the effects on entry difficult to predict. Spearot (2012) uses a model with variable markups to study the impact of tariff declines, but his theoretical results abstract away from the extensive margin response.

7. Conclusion

We used the Chinese experience following its entry into the WTO in 2001 to verify how the response of the extensive import margin to a tariff decline varies across different types of products. We use an extremely flexible estimation approach to control for comparative advantage, bilateral trade costs, and disaggregated demand and supply shocks. The tariff elasticities are identified solely from “within” changes for country-product observations and the export response to China is further normalized by the export experiences on the Indian and Indonesian markets.

Three findings are worth highlighting. First, the extensive margin response is relatively strong, precisely estimated, and robust to variations in the estimation specification. The absolute magnitude of the response to the tariff cuts accounts for a sizable fraction of the overall change in import penetration over the sample period. Second, the response is always estimated to be higher for differentiated goods or for goods with a low import demand elasticity. This pattern fits the theoretical prediction of Besedeš and Cole (2017) which shows that the tariff elasticity at the extensive margin should differ from a variable cost elasticity. Third, the responses are much stronger for OECD than for non-OECD countries, but reassuringly they are completely absent for non-WTO members.

While the absolute size of the tariff elasticity is always estimated higher for OECD countries, tariff cuts often account for a larger fraction of the observed increase in import penetration for non-OECD countries. This holds in most cases, but especially for differentiated products and for products with a low import demand elasticity, product categories with a strong response everywhere, It illustrates that trade liberalization can be an important source of long-run trade growth for developing countries. Given their current relatively low levels of import penetration, extensive margin responses are especially important for them.

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Table A.1: List of exporters

Group	Countries
OECD countries (WTO members) (33 countries)	Australia, Austria, Belgium/Luxemburg, Canada, Chile, Czech Rep., Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, India, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Republic of Korea, Slovakia, Slovenia, Spain, Sweden, Switzerland, Turkey, United Kingdom, United States
Non-OECD countries (WTO members) (91 countries)	Albania, Angola, Antigua and Barbuda, Argentina, Bahrain, Bangladesh, Barbados, Belize, Benin, Bolivia, Botswana, Brazil, Brunei, Bulgaria, Burkina Faso, Burundi, Cameroon, Chad, Colombia, Congo, Costa Rica, Croatia, Cuba, Cyprus, Djibouti, Dominica, Dominican Republic, Ecuador, Egypt, El Salvador, Fiji, Gabon, Georgia, Ghana, Grenada, Guatemala, Guinea, Guyana, Haiti, Honduras, Indonesia, Jamaica, Jordan, Kenya, Kuwait, Kyrgyzstan, Latvia, Lesotho, Liechtenstein, Lithuania, Madagascar, Malawi, Malaysia, Maldives, Mali, Malta, Mauritania, Mauritius, Central Africa, Moldova, Mongolia, Morocco, Mozambique, Myanmar, Namibia, Nepal, Nicaragua, Niger, Nigeria, Oman, Pakistan, Panama, Papua New Guinea, Paraguay, Peru, Philippines, Qatar, Cote d'Ivoire, Romania, Rwanda, Saint Lucia, Saudi Arabia, Senegal, Sierra Leone, Singapore, Solomon Islands, South Africa, Sri Lanka, Surinam, Swaziland, Tanzania, Thailand, Togo, Trinidad and Tobago, Tunisia, Uganda, United Arab Emirates, Uruguay, Venezuela, Zambia, Zimbabwe
Non-WTO members (11 countries)	Algeria, Equatorial Guinea, Ethiopia, Greenland, Iraq, Liberia, Dem. People's Rep. of Korea, Russian Federation, Sudan, Turkmenistan, Uzbekistan

Table A.2: Results for discrete product categories based on demand elasticities

	Dependent variable is a dummy for positive exports		
	All countries (1)	OECD countries (2)	Non-OECD countries (3)
(a) All products			
ln(1+*tariff*)			
* Low elasticity category	-0.123***	-0.243***	-0.079***
	(0.006)	(0.031)	(0.007)
* Medium elasticity category	-0.028***	-0.037	-0.024***
	(0.007)	(0.031)	(0.009)
* High elasticity category	-0.047***	-0.117***	-0.021**
	(0.007)	(0.027)	(0.009)
Observations	2,473,800	658,350	1,815,450
(b) Only differentiated products			
ln(1+*tariff*)			
* Low elasticity category	-0.177***	-0.352***	-0.113***
	(0.011)	(0.037)	(0.011)
* Medium elasticity category	-0.046***	-0.120**	-0.020***
	(0.017)	(0.057)	(0.005)
* High elasticity category	-0.019	-0.137*	0.024
	(0.016)	(0.076)	(0.017)
Observations	1,698,552	452,034	1,246,518

Notes: Product category dummies that interact the tariff variable are defined using the elasticity of China's import demand as estimated by Broda et al. (2006). Products are divided equally in low, medium, and high elasticity groups. Same specification as in the benchmark case: sample period is 2001 and 2006, export destinations are China, India, and Indonesia. In panel (b), differentiated products are identified using the conservative classification of Rauch (1999). ***, **, and * indicate significance at the 1%, 5% and 10% level.