Exporters and Shocks*

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Abstract

We use micro data for Ireland to estimate the responses of export entry, export exit, and the export revenue of incumbent exporters to changes in tariffs and real exchange rates. Entry and revenue are much more responsive to tariffs than they are to real exchange rates. Our estimates translate into an elasticity of aggregate exports with respect to tariff changes of between -1.5 and -3.5 on impact, and between -2 and -5 in the long run. Comparable elasticities for real exchange rate changes are around 0.5 on impact, and between 0.6 and 0.8 in the long run. These estimates are consistent with estimates in the literature based on aggregate data. They provide further evidence that workhorse models of international trade and business cycles which impose identical responses must be modified in order to answer policy questions touching on both international trade and the current account.

Keywords: Exporters; Tariffs; Real Exchange Rates; International Elasticity Puzzle

1 Introduction

We use customs and product-level production data for Ireland to estimate the responses of export entry, exit, and the export revenue of incumbent exporters to changes in ad valorem

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tariffs and real exchange rates. The direction of responses is as predicted by theory: lower tariffs and depreciations of the domestic real exchange rate both lead to increases in export entry and higher export revenue for incumbents. But export participation and export revenue are much more responsive to changes in tariffs than to comparable changes in real exchange rates: export entry is three times more responsive, while export revenue is nearly six times more responsive.

Our empirical strategy for estimating these responses is reduced form, though informed by models of the exporter problem structurally estimated in the literature. We use linear probability models to estimate the average marginal effects of variation in tariffs and real exchange rates on export entry and exit. We separately estimate elasticities of export revenue with respect to these variables. We deal with possible selection bias by focusing on the responses of observations which, based on their export history, have a very high probability of export participation. In all cases, we estimate responses to log deviations in tariffs and real exchange rates from their averages over the sample at the product-market level, thus using time-series variation to identify the coefficients of interest. Moreover, we control for marginal cost by comparing behavior of the same firm-product pair in markets where they face different shocks. In addition, we control throughout for real aggregate demand in the relevant market.

We use our micro estimates of responses to tariffs and real exchange rate to calculate implied elasticities of aggregate exports with respect to these shocks. In doing so, we take account of systematic patterns of post-entry selection and dynamics in export revenue and export exit. Our point estimates of micro responses translate into an elasticity of aggregate exports with respect to tariff changes of between -1.5 and -3.5 on impact, and between -2 and -5 in the long run. Elasticities with respect to changes in real exchange rates are around 0.5 on impact, and between 0.6 and 0.8 in the long run. These elasticities are comparable to estimates in the literature of elasticities of aggregate exports with respect to these variables.[1]

We are the first to provide an integrated empirical treatment of participation and revenue responses to both tariffs and real exchange rates, and to show how these responses can be aggregated to generate elasticities of aggregate exports. However micro-level responses to these two variables have been previously explored by two separate lines of research. Campa (2004) estimates the responses of export participation and export revenue to real exchange rates. Our findings are similar to his. They are also similar to those of Berman, Martin and Mayer (2012) who use French data to estimate participation, quantity, and price elasticities

[1]See Ruhl (2008) for a summary of the evidence.

to exchange rates. In the trade literature, most empirical work estimating micro responses to tariffs focuses on export participation, and uses cross-firm rather than time-series variation to identify responses: see e.g. Lileeva and Trefler (2010) and Bustos (2011). In work that is subsequent to ours, Fontagné, Martin and Orefice (2017) use customs data to estimate export revenue (but not participation) responses to tariffs and real exchange rates. Their findings are comparable to ours.

Understanding how exporters respond to tariffs and real exchange rates is of interest for two reasons. First, in workhorse models of international trade and business cycles, export responses to tariff reductions are identical to the expenditure-switching effects of real exchange rate movements. But in contrast to what these models predict, aggregate exports are systematically found to be more responsive to trade liberalizations than they are to movements in international relative prices arising from exchange rates. This is a challenge for quantitative analysis of any question touching simultaneously on international trade and the behavior of the current account, and begs for guidance as to how these models should be modified.

Second, a number of important policy questions do touch on both international trade and the current account, and answers to these questions hence depend on how exports respond to both changes in tariffs and exchange rates. These include whether "currency manipulation" is equivalent to trade protection,[2] and whether a border adjusted corporation tax will be offset by nominal exchange rate appreciation, or will have an impact on international trade.[3] More generally, the equivalence between trade policy and alternative forms of government intervention depends on how firms respond to shocks to relative prices arising from different sources. With the caveat that the exchange rate responses we estimate are identified from all sources of exchange rate variation, not just variation associated with tax changes or currency manipulation, our findings suggest that combinations of alternative policies that affect firms through real exchange rates should not be viewed as equivalent to changes in trade policy.

Meanwhile, our results provide some guidance as to how workhorse models might be modified to better reflect aggregate responses to shocks, and hence be better suited to policy analysis. Some of this guidance is negative. We show that the behavior of aggregate exports is not solely an artifact of aggregation,[4] since both export entry and the export revenue of individual exporters respond differently to tariffs and real exchange rates. In addition, it cannot be, as is sometimes argued, that the sole reason why aggregate exports appear to

[2]See Staiger and Sykes (2010) and Gagnon (2012) for contrasting positions on this issue.
[3]See Barbiero, Farhi, Gopinath and Itskhoki (2017).
[4]See Imbs and Mejean (2015) for this argument.

respond differently to the two shocks is that long run responses to tariffs are inappropriately compared with short run responses to real exchange rates. Our micro-level responses are identified using only time-series variation over the same time horizon for both tariffs and real exchange rates, yet we find very different responses.

On a more positive note, our findings are consistent with recent research which suggests at least two factors which may contribute to the greater responsiveness of exports to tariffs than to real exchange rates. Ruhl (2008) points out that irreversibilities in export decisions (e.g. sunk costs of export participation) may interact with the different statistical properties of tariffs (persistent, not very volatile) and real exchange rates (less persistent, more volatile) to generate different participation responses to the two shocks. Meanwhile Fitzgerald, Yedid-Levi and Haller (2017) show that when expenditures on marketing and advertising take place in the destination market, as in Arkolakis (2010) and Drozd and Nosal (2012), export revenue responses to real exchange rates may be muted. This does not affect responses to tariffs, potentially explaining differences in the responsiveness of export revenue to the two shocks.

The paper is laid out as follows. The second section describes our micro data. The third section describes our tariff and macro data. The fourth section motivates and describes our empirical strategy. The fifth section describes our main results. In the sixth section, we use estimates of post-entry export dynamics to aggregate from individual exporter responses to the response of aggregate exports to tariffs and real exchange rates. The final section of the paper concludes.

2 Micro data

We make use of three sources of confidential micro data made available to us by the Central Statistics Office (CSO) in Ireland: the Census of Industrial Production (CIP), customs records, and the Prodcom survey.

2.1 Census of Industrial Production

The CIP is an annual census of firms and establishments in manufacturing, mining and utilities. Firms with 3 or more persons engaged are required to file returns.[5] Because customs data is matched to firms, not plants, we focus on the firm as the unit of analysis. We use these data to construct the set of potential export participants for our analysis of

[5]Multi-plant firms also fill in returns at the level of individual plants, but we work with the firm-level data since this is the level at which the match with the customs data can be performed.

export participation, and also to condition on firm characteristics. We use data for the years 1996 through 2009, and for NACE Revision 1.1 sectors 10-40 (manufacturing, mining and utilities). Of the variables collected in the CIP, those relevant for our purposes are total revenue and employment.

In constructing the sample for analysis, we drop firms with a zero value for total revenue or zero employees in more than half of the years they are present in the data. We perform some recoding of firm identifiers to maintain the panel dimension of the data, e.g. in cases where ownership changes. Further details on the data and how we have cleaned them are provided in the Appendix.

2.2 Customs records

Our second source of data is customs records of Irish merchandise exports for the years 1996 through 2009. These records consist of the value (in Euros) and volume (in tonnes) of exports at the level of the VAT number, the Combined Nomenclature (CN) 8-digit product, and the export market (country). We use only the value data. These data are aggregated by the CSO to an annual frequency, and then matched by them to CIP firms using a correspondence between VAT numbers and CIP firm identifiers, along with other confidential information. In the Appendix, we provide summary statistics on this match. We classify firms as exporters if they are matched with positive exports from the customs data. Match quality appears to be poor in 1996, so in our baseline empirical analysis we assume that export participation is censored in 1996.

An important feature of the customs data is that the 8-digit CN classification system changes every year. We concord the product-level export data over time at the most disaggregated level possible following the approach of Pierce and Schott (2012), as implemented by van Beveren, Bernard and Vandenbussche (2012). We make use of these concorded data in our analysis of export revenue. For our analysis of export participation we further concord the export data with production data, as described below.

Another important feature of customs in the EU is that data for intra-European and extra-European trade are collected separately, using two different systems, called Intrastat and Extrastat. For Ireland, the reporting threshold for intra-European exports (635,000 Euro per year in total shipments within the EU)[6] is much higher than the reporting threshold for extra-European exports (254 Euro per transaction). However the intra-European threshold

[6]Intra-European exports below the threshold are recovered based on VAT returns. The destination within the EU is not recorded for these returns. Since we do not know the exact destination, we do not make use of these exports other than in our summary statistics on exporters vs. non-exporters.

applies to *total* intra-European exports, not at the individual country level, and it is common for firms to report exports to individual EU countries which are well below the 635,000 Euro threshold. In our baseline analysis we pool both Intrastat and Extrastat exports, ignoring the fact that export participation may be censored for Intrastat markets. We check that our results are robust to restricting the sample to Extrastat markets, and to export to Intrastat markets by firms with total Intrastat exports above the reporting threshold.

2.3 Prodcom survey

Our third source of data is the Prodcom survey for the years 1996 to 2009. This is an annual survey of the value (in Euros) and volume (in tonnes) of all products manufactured by the firm and sold in the relevant year. We use only the value data. The survey basis is all firms in the CIP, excluding some mining sectors. On average over the sample, just under 65% of CIP firms, accounting for just under 95% of total CIP revenue participate in the Prodcom survey. We use these data to define the set of products a firm might potentially export.

Products are classified at the 8-digit level according to the Prodcom classification. As with the CN 8-digit classification, the 8-digit Prodcom classification system changes from year to year. We first concord the Prodcom data over time at the most disaggregated level possible. We then concord the time-concorded Prodcom data to the time-concorded CN data. In doing so, we follow van Beveren, Bernard and Vandenbussche (2012). There are fewer Prodcom 8-digit products than there are CN 8-digit products, so this involves some aggregation at the CN8 level.

As a baseline, we define the set of products a firm might potentially export in any given year as all products that it produces at least once during its lifetime in the CIP. If the firm does not export one of these products to a particular destination market in a particular year, we code this as non-participation at the firm-product-market-year level. We also check robustness to defining the set of products that a firm might potentially export in a given year as the products it actually produces in that year.

2.4 Summary statistics

Table 1 presents summary statistics on firms and exports. Firms matched to exports in the customs data are on average bigger than firms which sell exclusively to the domestic market, though there is substantial overlap in the size distributions, as illustrated in Figure 1. Export participation and export intensity for exporters are both relatively high, as is

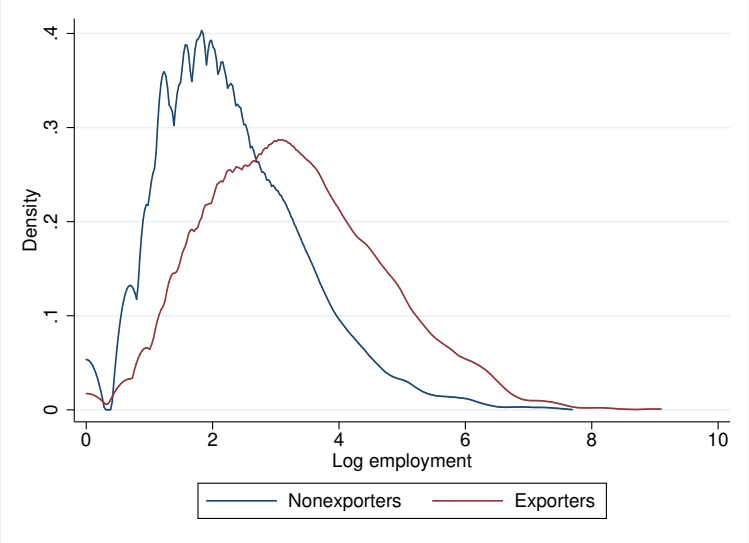
typical of small open economies in Europe (see ISGEP (2008)). Export participation is quite persistent, a fact which is again typical of other countries for which micro data on exporting are available.

Table 1: Summary statistics: Firms and exports, averages 1996-2009

Number of firms per year	4729
Employees per firm	51
Firm age (years)	17
Share of firms foreign owned	0.13
Share of multi-plant firms	0.03
Export participation rate	0.42
Export entry rate	0.09
Export exit rate	0.12
Exporter size premium (employees)	3.29
Exporter size premium (revenue)	5.20
Export share conditional on exporting	0.33
Number of markets per exporter	6.9

Notes: Statistics are for our cleaned data set of CIP firms. Firms are defined as exporters if they are matched to positive exports from customs data. Export intensity is calculated as total exports from customs divided by sales reported in the CIP. Values greater than 1 are replaced by 1. The set of potential entrants used to calculate the entry rate in year t includes firms born in year t. The set of potential exiters used to construct the exit rate in year t includes firms present in the CIP in year $t-1$ but not in year t. Entry and exit rate averages are calculated over 1998-2009. Source: CSO and authors' calculations.

Figure 1: Distribution of log employment for non-exporters and exporters



Notes: Statistics are for our cleaned data set of CIP firms. Firms are defined as exporters if they are matched to positive exports from customs data. Figure plots kernel density estimates of the distribution of log employment for non-exporters and exporters, pooling across 1996-2009. Source: CSO and authors' calculations.

In our analysis of responses to tariffs and real exchange rates, we restrict attention to 30 of Ireland's largest export markets (countries). The exact choice of markets is governed by tariff availability. These markets account for 94% of total exports in the sample period. Table

2 reports the breakdown of exports by market for the most important of these 30 markets. The main export markets for our firms are the US, the UK and the large economies of the Eurozone.

Table 2: Percentage of exports by destination

	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
Australia	1	1	1	1	1	1	1	1	1	1	1	1	1	1
Canada	0	0	0	0	0	1	0	1	0	0	0	0	0	1
China	0	0	0	0	0	0	1	1	1	1	1	2	2	2
Denmark	1	1	1	1	1	1	1	1	1	1	1	1	1	0
Japan	2	3	3	3	4	4	3	3	3	3	3	2	2	2
Norway	1	1	1	1	1	1	1	1	1	1	1	1	0	0
Sweden	2	2	2	2	2	2	1	1	1	1	1	1	1	1
Switzerland	2	2	2	2	3	3	4	4	4	5	4	4	4	4
UK	24	24	21	18	18	17	16	15	16	16	17	17	17	15
US	11	12	16	16	20	19	21	23	21	16	16	17	18	22
Euro 9	43	41	43	47	42	41	43	42	44	47	45	43	43	44
Other	5	4	3	4	4	5	4	4	4	3	4	4	5	4
Total in-sample	93	92	93	95	93	94	95	95	94	94	93	93	94	95

Notes: We use only customs export records which match to our cleaned data set of CIP firms. The Euro 9 includes Austria, Belgium, Finland, France, Germany, Italy, Netherlands, Portugal and Spain. Other includes Brazil, Hong Kong, India, Malaysia, Mexico, New Zealand, Saudi Arabia, South Africa, Thailand, Turkey and the United Arab Emirates. Source: CSO and authors' calculations.

We now restrict attention to these 30 export markets, and look at churn in export status at the level of the individual market. We define a firm-market observation as a potential entrant at date t if the firm is present in the CIP at date t, and did not export to that market in $t-1$.[7] We define a firm-market observation as a potential exiter at date t if it had positive exports in year $t-1$.[8] The definition for entry and exit at the firm-product-market is analogous, with the caveat that the firm must not only be present in the CIP, but also in Prodcom, in order for participation to be defined. Table 3 reports average entry, exit and participation rates at the firm-market and firm-product-market level based on these definitions. It also reports average entry, exit, and participation rates by firm size.

Table 3 illustrates three important facts. First, there is steady state churn in exporting at the firm-market and firm-product-market level. Second, export entry, exit, and participation vary systematically with firm size. Third, past export participation is a much stronger predictor of current export participation than is firm size.

History dependence in exporting goes beyond a strong relationship between current participation and participation in the previous year. Tables 4 and 5 illustrate this at the firm-product-market level. Potential entrants with past export experience (i.e. experience before the previous year) are much more likely to enter than those without such experience.

[7]Entering firms are categorized as potential entrants in all markets.

[8]Exiting firms are included.

Meanwhile, for incumbent exporters, exit rates decline dramatically (or conversely, survival rates go up) with tenure in the relevant market market. We both control for, and exploit these facts in our strategy for estimating responses to shocks.

Table 3: Average participation, entry, and exit rates, 1998-2009

Rate	All firms	Small†	Med†	Large†
	Firm-market level			
Entry	0.012	0.01	0.04	0.07
Exit	0.226	0.26	0.21	0.17
Particip.	0.051	0.03	0.18	0.29
	Firm-product-market level			
Entry	0.008	0.00	0.02	0.03
Exit	0.261	0.29	0.25	0.22
Particip.	0.028	0.01	0.07	0.12

Notes: Statistics are for our cleaned data set of CIP firms. Products are defined based on the concordance of Prodcom and CN product definitions as described in Sections 2.2 and 2.3. The set of potential entrants used to calculate the entry rate in year t includes firms and firm-products born in year t. The set of potential exiters used to construct the exit rate in year t includes firms present in the CIP in year $t-1$ but not in year t. † Small firms are those with fewer than 100 employees in the previous year. Medium firms are those with 100-249 employees in the previous year. Large firms are those with 250+ employees in the previous year. Source: CSO and authors' calculations.

Table 4: Export entry and past export participation

History	Entry rate	Share of potential entrants	Share of entrants
Last participated at $t-2$	0.227	0.005	0.151
last participated at $t-3$	0.139	0.003	0.061
Last participated before $t-3$	0.073	0.007	0.064
Never participated	0.006	0.985	0.724

Notes: Statistics are for our cleaned data set of CIP firms. Products are defined based on the concordance of Prodcom and CN product definitions as described in Sections 2.2 and 2.3. The set of potential entrants used to calculate the entry rate in year t includes all firm-product-markets present in year t which did not export in year $t-1$. This includes firms born in year t. Source: CSO and authors' calculations.

Table 5: Export exit and past export participation

Tenure	Exit rate	Share in export observations	Share of exiters
1 year	0.45	0.24	0.42
2 years	0.33	0.12	0.16
3 years	0.22	0.08	0.06
4 years	0.23	0.05	0.04
5 years	0.18	0.03	0.02
6 years	0.15	0.02	0.01
7+ years	0.14	0.03	0.02
censored	0.16	0.43	0.26

Notes: Statistics are for our cleaned data set of CIP firms. Products are defined based on the concordance of Prodcom and CN product definitions as described in Sections 2.2 and 2.3. The set of potential exiters used to construct the exit rate in year $t$includes all firm-product-markets exporting in year $t-1$, including firms no longer present in the CIP in year t. Tenure is the number of years a firm-product-market observation has been continuously exporting since entry. For observations where exporting is ongoing in 1996 or starts in 1997 we categorize tenure as censored by the start of the sample. Source: CSO and authors' calculations.

3 Tariff and real exchange rate data

3.1 Tariff data

Our analysis requires data on the tariffs faced by Irish firms in export markets. Irish firms do not face tariffs when exporting to EU countries. When the EU has a preferential trade agreement, this governs the tariffs Irish exporters face. In countries where the EU does not have a preferential trade agreement, they face the Most Favored Nation (MFN) tariff. Since the EU is a customs union, in all of the countries where Irish exporters do face tariffs, the actual level of tariffs is determined by bargaining between the EU and the relevant counterparties. Given that Ireland is one small member of the EU, this makes us fairly confident that reverse causality is not a problem in estimating export participation and revenue responses to destination market tariffs.

The EU has preferential trade agreements with Switzerland, Norway, Turkey, South Africa and Mexico for some or all of our sample period. We obtain information on the relevant preferential tariffs from the WTO and other sources (detailed in the Appendix). We obtain MFN tariff data from the WTO for the following countries: Australia, Brazil, Canada, China, Hong Kong, India, Japan, Malaysia, Mexico, New Zealand, Saudi Arabia, South Africa, Thailand, United Arab Emirates and the US. For several of the countries in which Irish exporters face tariffs, especially developing countries, tariff data are not available for all of the sample years. Under limited circumstances we interpolate tariffs for the years for which data are not available. Full details of the sources, construction and coverage of the tariff data are in the Appendix.

Tariff data are reported by the WTO using the Harmonized System (HS) 6-digit classification. This is used by all countries as the basis for their tariff lines, and is the most disaggregated level at which tariffs, export, and production data can be matched across countries. We restrict attention to HS6 product-market-years for which there are no non-ad-valorem tariffs,[9] and for which there is no sub-HS6 variation in ad valorem tariffs.[10] The HS6 classification changes in 2002 and 2007. We concord the classification over the period 1996-2009 following the approach of Pierce and Schott (2012). When this implies joining multiple HS6 categories together in a given period, we take the simple average of tariffs to construct tariffs at the concorded product-market-year level.

[9] Unlike ad-valorem tariffs, non-ad-valorem tariffs affect incentives to export differently depending on the firm's export price.

[10] We test robustness to relaxing this requirement.

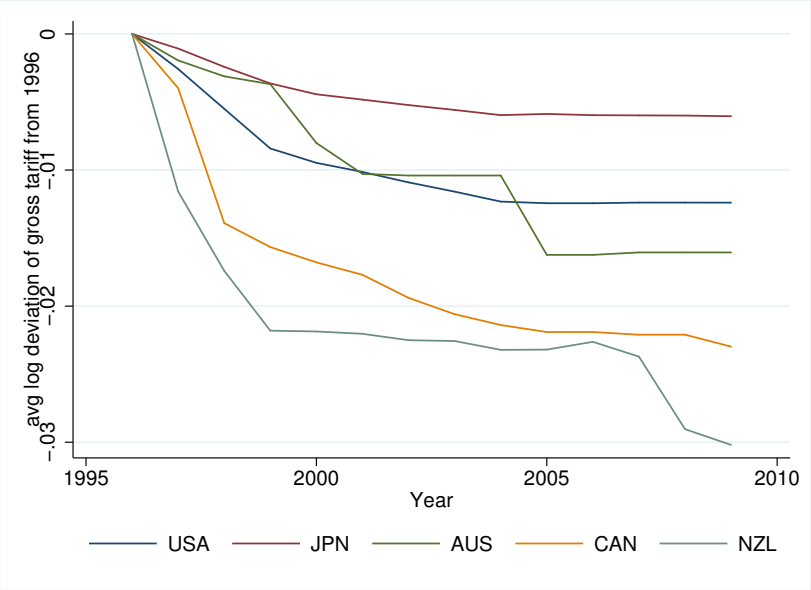


Figure 2: Average evolution of tariffs for five rich destination markets

Notes: Figure shows coefficients on year dummies in country-by-country regression of $\ln\left(1+T_t^{jk}\right)$ on HS6 fixed effects and year dummies. Source: WTO and authors' calculations.

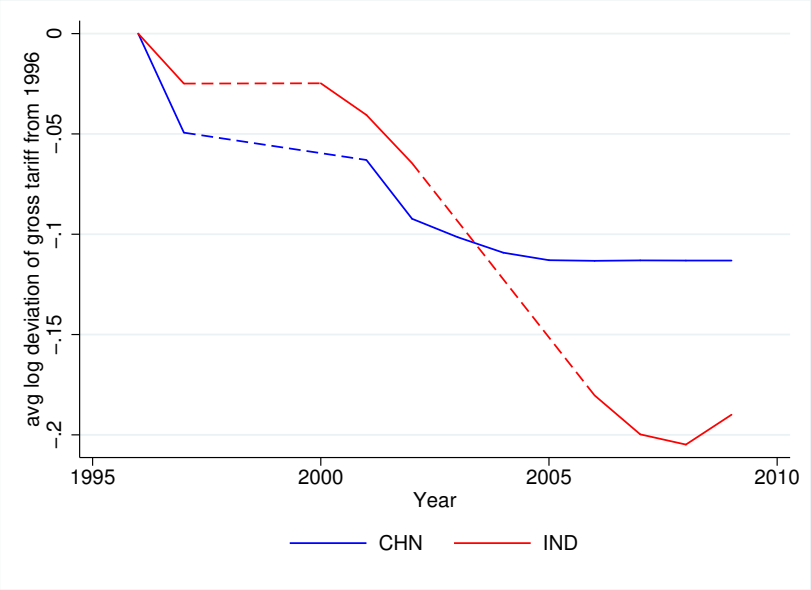


Figure 3: Average evolution of tariffs for China and India

Notes: Figure shows coefficients on year dummies in country-by-country regression of $\ln\left(1+T_t^{jk}\right)$ on HS6 fixed effects and year dummies. Dotted lines indicate missing years in data. Source: WTO and authors' calculations.

To make use of the tariff data, we must also concord it with our export and production data. At a 6-digit level, the CN (export) and Prodcom (production) classifications corre-

spond to the HS classification. In some cases, our concordance of the CN and Prodcom classifications results in "products" that cover multiple HS6 categories. In our analysis of export entry and exit, in these cases we take the simple average of tariffs across the relevant HS6 categories to obtain a tariff at the concorded product-market-year level. In our analysis of export revenue, we use revenue at the firm-HS6-market-year level to construct a weighted average of tariffs across the relevant HS6 categories.[11]

In our baseline analysis we rely on time-series variation in tariffs to identify exporter responses.[12] The main source of time-series variation in tariffs faced by Irish exporters in our sample is the gradual phasing-in over the period 1996-2005 of MFN tariff reductions agreed in the context of the Uruguay Round of the WTO in 1994. To illustrate the nature of this variation, we regress the log gross ad valorem tariff, i.e. $\ln\left(1+T_t^{jk}\right)$, on HS6 fixed effects and year dummies, market-by-market, for a selection of important markets. The coefficients on the year dummies for five rich tariff-subject markets are plotted in Figure 2, while Figure 3 plots the year dummies for China and India. These figures illustrate the fact that average year-on-year changes in tariffs for many of the large tariff-subject markets Ireland trades with is small. What the figures do not show is that for many markets, tariff changes are concentrated in a narrow set of products. This presents a challenge for identifying responses to tariff changes. Additional summary statistics on the tariff data are reported in the Appendix.

3.2 Macro data

Our analysis requires measures of the real consumption exchange rate between Ireland and the 30 destination markets of interest. We also require controls for real aggregate demand. We construct indexes for real exchange rates using data on annual average nominal exchange rates and CPIs from the IMF's *International Financial Statistics* (IFS). We construct an index of real demand using GDP less exports plus imports, all measured in current local currency, with this aggregate deflated by the relevant country's CPI. National Accounts data are from the OECD's *National Accounts Statistics* where available, and otherwise from the World Bank's *World Development Indicators*. CPIs are from IFS.

The bulk of the variation in real exchange rates is driven by nominal exchange rates. This is illustrated in Figure 4, which contrasts the evolution of log deviations of real exchange

[11] We also check that results are robust to this choice by dropping the affected cases.

[12] We want to use similar sources of variation to identify responses to tariffs and real exchange rates. Since we do not measure cross-sectional variation in real exchange rates, we cannot make use of cross-sectional variation in tariffs.

rates vis-a-vis selected non-Euro markets from their 1996 levels with the evolution of real exchange rates vis-a-vis the nine Eurozone markets in our sample. There is more year-on-year variation in real exchange rates for non-Eurozone markets than there is in tariffs.

For most of the sample period (i.e. 1998-2009), monetary policy is decided at the Eurozone level by the ECB, mitigating concerns about reverse causality. In a robustness check, we exclude markets which use the Euro so as to further mitigate concerns about reverse causality.

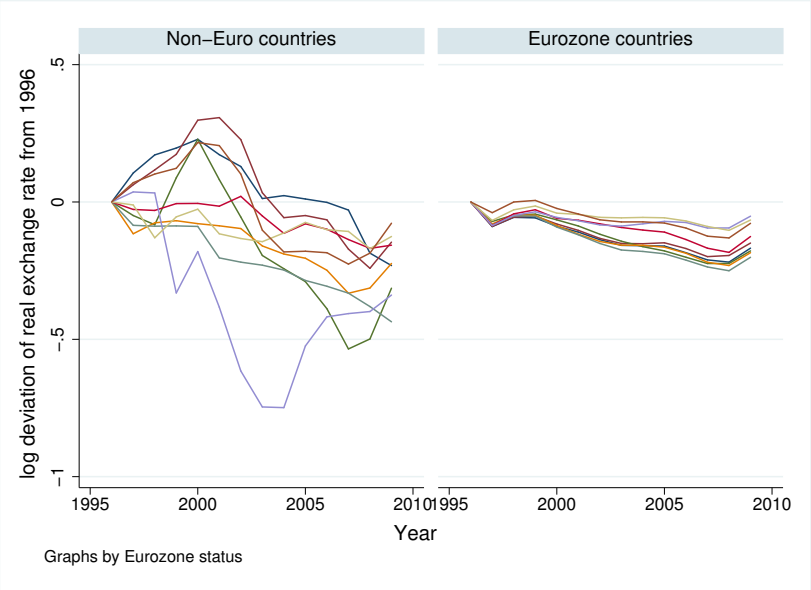


Figure 4: Evolution of real exchange rates in selected non-Euro and Eurozone markets

Notes: Figure shows log deviation of annual average real exchange rate from 1996 level for selected Non-Euro and Eurozone countries. The Non-Euro countries are Australia, Brazil, China, Japan, Norway, Sweden, Switzerland, UK and USA. The Eurozone countries are Austria, Belgium, Finland, France, Germany, Italy, Netherlands, Spain and Portugal. Source: IMF and authors' calculations.

4 Empirical strategy

The goal of our empirical analysis is to estimate how sensitive export participation and export revenue are to changes in tariffs and real exchange rates. We adopt a reduced form strategy, which has the advantage that our estimates can thus be used to discipline a variety of structural models. Nevertheless, we motivate our estimation strategy by first sketching a model of the exporter's decision problem. This nests a variety of different models structurally estimated in the literature.[13]

[13]See, e.g. Das, Roberts, and Tybout (2007), Ruhl & Willis (2016), Fitzgerald, Haller, and Yedid-Levi (2017).

Firm i's net profit flow from exporting product j to country k at date t is equal to export revenue, less total associated costs. This net profit flow is a function of date-t actions taken by the firm (A_t^{ijk}), and firm-level (Y_t^{ijk}) and aggregate (Z_t^{jk}) state variables. Actions include whether the firm exports in period t ($X_t^{ijk} \in \{0,1\}$), and if so, what price or quantity it sets. Actions may also include marketing and advertising expenditures designed to shift demand in market k. The firm-level state variables include (a) the exporter's marginal cost of producing good j, along with stochastic fixed or sunk costs of export participation, (b) idiosyncratic demand, and (c) the firm's history of actions in market k. This history may affect net profits through costs (e.g. entrants pay sunk costs, but incumbents do not) or through demand (e.g. more advertising in the past implies higher demand today). Aggregate state variables may include (d) the tariff incurred by importers of the firm's product in country k, (e) the real exchange rate between the home market and market k, and (f) real aggregate demand in market k.[14]

All of the exogenous state variables, both firm level and aggregate, evolve over time following processes known to the firm. The firm's history of actions evolves consistent with A_t^{ijk}. Assuming that firms discount the future at rate β, the Bellman equation for the firm's problem can be written:

$$V\left(Y_t^{ijk}, Z_t^{jk}\right) = \max_{A_t^{ijk}} \mathbb{E}\left\{\Pi\left(A_t^{ijk}, Y_t^{ijk}, Z_t^{jk}\right) + \beta V\left(Y_{t+1}^{ijk}, Z_{t+1}^{jk}\right)\right\},$$

subject to the evolution of firm-level endogenous state variables as a function of actions A_t^{ijk}.

This gives rise to a policy function:

$$A_t^{ijk} = A\left(Y_t^{ijk}, Z_t^{jk}\right)$$

In particular, it implies that the probability of participation today can be expressed as a function of the firm-level and aggregate state variables:

$$\Pr\left[X_t^{ijk} = 1\right] = \Pr\left[X\left(Y_t^{ijk}, Z_t^{jk}\right) > 0\right] \tag{1}$$

For those firms that do export, maximized export revenue can also be expressed as a function of firm-level and aggregate state variables:

$$R_t^{ijk} = R\left(Y_t^{ijk}, Z_t^{jk}\right) \tag{2}$$

[14]This list of actions and state variables need not be exhaustive.

In some popular models, export revenue is a function of a strict *subset* of firm-level state variables. For example, in Das, Roberts and Tybout (2007), conditional on the persistent component of unobserved idiosyncratic demand, revenue is independent of the history of export participation. More generally, e.g. as in Ruhl and Willis (2016) and Fitzgerald, Haller, and Yedid-Levi (2017), export revenue does depend on the history of actions even conditional on the persistent component of idiosyncratic demand, and we allow for this.

4.1 Export entry and exit

We use a reduced form approximation to equation (1) to directly estimate the sensitivity of export participation to tariffs and real exchange rates. Our baseline specification is a linear probability model, so coefficient estimates can be interpreted as average marginal effects. Motivated by the model, and by the stylized facts about export participation documented in Section 2, we control for time-varying heterogeneity in marginal cost of production, and for export histories, in addition to the shocks of interest.

Tariffs, real exchange rates and foreign demand are included in our regressions in logs, along with product-market fixed effects. Since tariffs vary at the product-market-year level, and real exchange rates and foreign demand at the market-year level, this means that the coefficients on these variables can be interpreted as the average marginal effects on entry and exit of log deviations of these variables from their average levels over the sample for the relevant product-market.

We control for marginal cost of production using firm-product-year fixed effects. As a result, we identify responses to shocks based on differential behavior of the same firm-product in two different markets where the deviation of shocks from their average level is different.

We control for export histories using a rich set of indicator variables. In the entry equation, we include indicator variables for last participation at date $t-2$ (remember that non-participation at $t-1$ classifies an observation as a potential entrant), for last participation at date $t-3$, and for last participation before date $t-3$. The omitted category is observations where we do not observe any past participation in-sample. In the exit equation, we include indicator variables for the number of years since the most recent export entry, topcoding at 7. We also include an indicator for censored entry (i.e. censored by the beginning of the sample).

It should be noted that the export history variables capture not just any causal relationship between export histories and entry and exit probabilities (e.g. due to sunk costs of participation), but also the impact of persistent heterogeneity in idiosyncratic demand.

This is because observations with a strong persistent component of idiosyncratic demand are likely to have a long history of export participation. Since we are not directly interested in whether export persistence is due to persistent unobserved heterogeneity or frictions such as sunk costs, it is not necessary to separate out these two sources of persistence.

The major disadvantage of this specification relative to e.g. a conditional logit or random effects probit model is that marginal effects are not allowed to differ across the distribution of independent variables. Because the average probability of export participation differs so strikingly for potential entrants and incumbents (i.e. 0.01 vs 0.84), it is highly likely that marginal effects will be different for these two groups. So we estimate separate equations for entry and exit rather than a single equation for export participation. In robustness checks, we also include interaction terms in our entry and exit equations in order to to estimate different average marginal effects for different subsets of potential entrants and incumbents.

Our baseline specifications for entry and exit are:

$$\Pr\left[X_t^{ijk}=1|X_{t-1}^{ijk}=0\right]=c_t^{ij}+\gamma^{jk}+\boldsymbol{\lambda}'\mathbf{a}_{t-1}^{ijk}+\boldsymbol{\beta}'\mathbf{z}_t^{ijk}+\eta_t^{ijk} \tag{3}$$

$$\Pr\left[X_t^{ijk}=0|X_{t-1}^{ijk}=1\right]=c_t^{ij}+\gamma^{jk}+\boldsymbol{\lambda}'\mathbf{a}_{t-1}^{ijk}+\boldsymbol{\beta}'\mathbf{z}_t^{ijk}+\eta_t^{ijk} \tag{4}$$

X_t^{ijk} is, as above, the indicator for participation, c_t^{ij} is a firm-product-year fixed effect, γ^{jk} is a product-market fixed effect, and $\mathbf{a}_{t-1}^{ijk}$ is the vector of indicator variables for export history. $\mathbf{z}_t^{ik}$ is a vector, the elements of which are $\tau_t^{jk}=\ln\left(1+T_t^{jk}\right)$, which is the log gross ad valorem tariff faced by exporter i in market k, rer_t^k, which is the log of an index of the real exchange rate between the home market and market k, and dem_t^k, which is the log of an index of real aggregate demand in market k. The construction of tariffs and macro variables is described above in sections 3.2 and 3.1. We use robust standard errors to deal with the standard heteroskedasticity issue in estimating linear probability models.

4.2 Revenue

For export revenue, we approximate equation (2) with a log-linear equation.[15] As with export entry and exit, we include product-market fixed effects. The coefficients on tariffs and macro variables can thus be interpreted as the elasticity of export revenue with respect to deviations of these variables from their average levels over the sample for the relevant product-market.

[15]If demand is CES, revenue is log-linear in marginal cost, tariffs, real exchange rates, and foreign demand.

We control for time-varying heterogeneity in the marginal cost of production using firm-product-year fixed effects. The elasticities of interest are therefore identified by differential behavior of revenue of the same firm-product in two different markets where the deviation of shocks from their average level is different.

We control for export histories by including a set of indicator variables for the number of years since the most recent export entry (topcoded at 7 years), as well as an indicator for whether export entry is censored by the beginning of the sample. As in the case of entry and exit, the export history variables capture not just the causal relationship between export histories and revenue, but also the impact of persistent heterogeneity in idiosyncratic demand on revenue.

A key issue in estimating the elasticity of export revenue with respect to shocks is the possibility of selection bias. A shock to tariffs or real exchange rates that makes export participation more (less) attractive implies higher (lower) participation of firm-product-market observations with weak unobserved idiosyncratic demand. This induces correlation between the variables of interest and the error term, and hence coefficient estimates are likely to be biased. The standard econometric approach in the presence of selection bias in cross-section data is to use a Heckman selection correction. But we do not have any variables which plausibly affect export participation, but not export revenue conditional on participation. As a result, a Heckman-style selection correction would rely on functional form alone for identification.[16]

So we take an alternative approach based on our particular environment. First, note that conditional on export histories, selection bias is due to non-random truncation of the *transitory* component of unobserved idiosyncratic demand.[17] Next, note that high (low) realizations of the transitory component of idiosyncratic demand are less likely to induce entry (exit) for observations far from entry and exit thresholds than for observations close to those thresholds.[18] Given the strong relationship between export histories and participation, we can identify observations far from the participation thresholds based on these histories.

We implement this by including all observations in the regression, but allowing elasticities with respect to tariffs and macro variables to differ between observations with long export

[16]Somewhat less crucially, other dimensions of our identification strategy rely on a rich set of fixed effects which are tricky to incorporate in the standard Heckman setup.

[17]In any case, selection on the persistent component of idiosyncratic demand is less of a concern, because we are not exploiting cross-sectional variation in tariffs or real exchange rates to identify responses.

[18]Loosely speaking, this idea resembles Chamberlain's (1986) "identification at infinity" approach to dealing with selection bias. Mulligan and Rubinstein (2008) provide an implementation of the method in a different context.

histories and those with short export histories. In the long history group, we include observations with 6+ years of tenure, and observations where entry is censored by the beginning of the sample. These observations have low exit rates (around 15% anually),[19] and account for 48% of observations, but 75% of in-sample exports. The short history group consists of observations with 1-5 years of tenure. The coefficients for observations in the long export history group are less likely to be subject to selection bias.

Our baseline estimating equation for export revenue conditional on participation is then:

$$r_t^{ijk} = c_t^{ij} + \gamma^{jk} + \boldsymbol{\lambda}' \mathbf{a}_t^{ijk} + \boldsymbol{\beta}' \left(\mathbf{z}_t^{ik} * long_t^{ijk} \right) + \boldsymbol{\phi}' \left(\mathbf{z}_t^{ik} * short_t^{ijk} \right) + \varepsilon_t^{ijk} \tag{5}$$

Here, r_t^{ijk} is the log of export revenue of firm i from selling product j to market k at date t, c_t^{ij} is a firm-product-year fixed effect, and γ^{jk} is a product-market fixed effect. As in the case of entry and exit, $\mathbf{a}_t^{ijk}$ is a vector of indicator variables for export history. $\mathbf{z}_t^{ik}$ is as described above. Finally, $long_t^{ijk}$ (for long export tenure) is an indicator variable, set equal to 1 for observations where market tenure is at least 6 years as well as observations in export spells whose entry is censored, and equal to zero otherwise, while $short_t^{ijk}$ is equal to $1 - long_t^{ijk}$.[20] The coefficients of interest are $\boldsymbol{\beta}'$.

5 Results

5.1 Entry and exit

The first column of Table 6 reports the results from estimating our entry equation (3), while the second column reports the results from estimating our exit equation (4). We report only the coefficients on the shock variables. Full results, including coefficients on export history controls, are reported in the Appendix. As predicted by theory, the probability of entry is higher when tariffs for the relevant product in the destination market are low relative to their average over the sample, and when the real exchange rate vis-a-vis the destination market is weak relative to its average over the sample. Both coefficients are statistically significant at the 5% level. We do not find evidence of a statistically significant relationship between the rate of exit and tariffs or real exchange rates.[21]

[19] See Table 5.

[20] Since $long_t^{ijk}$ and $short_t^{ijk}$ are linear combinations of $\mathbf{a}_t^{ijk}$, their level effects are captured by the inclusion of $\mathbf{a}_t^{ijk}$ in the regression.

[21] Though the standard error is large, the coefficient on tariffs in the exit equation has the correct sign, and is of a similar order of magnitude when compared to the exit rate as the coefficient on tariffs in the

Table 6: Export entry and exit responses to shocks

	(1)		(2)	
	Entry		Exit	
	coeff	s.e.	coeff	s.e.
τ_t^{jk}	-0.007	(0.002)**	0.127	(0.116)
rer_t^k	0.002	(0.001)**	0.015	(0.023)
dem_t^k	0.005	(0.001)**	0.013	(0.024)
Export history controls	yes		yes	
Firm-prod-yr f.e.	yes		yes	
Prod-mkt f.e.	yes		yes	
N	2,383,762		70,189	
R^2	0.29		0.60	
R^2-adjusted	0.24		0.49	
	In-sample entry rate		In-sample exit rate	
	0.008		0.218	

Notes: Products are defined based on the concordance of Prodcom and CN product definitions as described in Sections 2.2 and 2.3. Sample period is 1998-2009. Only potential entrants are included in the entry regression. Only potential exiters are included in the exit regression. Dependent variable is an indicator for entry or exit at the firm-product-market-year level. Export history controls in column (1) include include indicator variables for last participation at date $t-2$, last participation at date $t-3$, and last participation before date $t-3$. Export history controls in column (2) include indicators for export tenure (topcoded at 7 years), and an indicator for censored export tenure. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

The absolute value of the coefficient on the tariff in the entry equation is 3 times bigger than the coefficient on the real exchange rate (the difference is statistically significant at the 10% level), so entry is more responsive to deviations in tariffs from their average values than to deviations of real exchange rates from their average values. To understand the economic significance of these estimates, remember that the average in-sample entry rate is 0.8%. So a reduction in the tariff from 10% to 0 induces an increase in the entry rate from 0.8%, to 0.87%, while a 10% depreciation of the home real exchange rate against the destination market induces an increase in the entry rate from 0.8% to 0.82%. In Section 6, we calculate the dynamic impact of this entry response on aggregate exports.

5.1.1 Heterogeneous marginal effects

As noted in Section 4.1, the sensitivity of entry and exit to shocks to tariffs and real exchange rates may vary across observations. By averaging marginal effects across all observations in our baseline estimating equation, we run the risk of mistakenly inferring insensitivity to shocks. To address this issue, we allow average marginal effects to differ across groups of observations by interacting the shock vector $\mathbf{z}_t^{ijk}$ with indicators for firm size, market characteristics, and export history in turn.

entry equation relative to the entry rate. In contrast, the coefficient on the real exchange rate in the exit equation is both economically and statistically insignificant.

Table 7: Export entry and exit: Marginal effects for firms of different size

$$\text{Entry: } \Pr\left[X_t^{ijk}=1|X_{t-1}^{ijk}=0\right]=c_t^{ij}+\gamma^{jk}+\boldsymbol{\lambda}'\mathbf{a}_{t-1}^{ijk}+\boldsymbol{\beta}'\left(\mathbf{z}_t^{ijk}\otimes\mathbf{size}_{t-1}^{i}\right)+\eta_t^{ijk}$$

$$\text{Exit: } \Pr\left[X_t^{ijk}=0|X_{t-1}^{ijk}=1\right]=c_t^{ij}+\gamma^{jk}+\boldsymbol{\lambda}'\mathbf{a}_{t-1}^{ijk}+\boldsymbol{\beta}'\left(\mathbf{z}_t^{ijk}\otimes\mathbf{size}_{t-1}^{i}\right)+\eta_t^{ijk}$$

	Small		Medium		Large	
	coeff	s.e.	coeff	s.e.	coeff	s.e.
	Entry					
τ_t^{jk}	0.005	(0.002)**	-0.054	(0.004)**	-0.079	(0.007)**
rer_t^k	0.002	(0.001)**	0.003	(0.001)**	0.004	(0.002)**
dem_t^k	0.004	(0.001)**	0.006	(0.002)**	0.007	(0.002)**
Entry rate	0.005		0.022		0.032	
Share of observations	0.840		0.096		0.063	
	Exit					
τ_t^{jk}	0.224	(0.157)	0.076	(0.176)	0.074	(0.137)
rer_t^k	0.010	(0.023)	0.012	(0.023)	0.012	(0.023)
dem_t^k	0.013	(0.025)	0.011	(0.025)	0.010	(0.025)
Exit rate	0.238		0.217		0.192	
Share of observations	0.405		0.283		0.311	

Notes: Marginal effects are based on estimation of the above equation, where $\mathbf{size}_t^i$ is an indicator for firm size. Small: <100 employees. Medium: 100-249 employees. Large: 250+ employees. Sample is as in Table 6. Dependent variable is an indicator for entry or exit at the firm-product-market-year level. Firm-product-market fixed effects, product-market fixed effects, and baseline controls for export history are included. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

Table 8: Export entry: Marginal effects for low vs high participation markets

$$\text{Entry: } \Pr\left[X_t^{ijk}=1|X_{t-1}^{ijk}=0\right]=c_t^{ij}+\gamma^{jk}+\boldsymbol{\lambda}'\mathbf{a}_{t-1}^{ijk}+\boldsymbol{\beta}'\left(\mathbf{z}_t^{ijk}\otimes\mathbf{mkt}^{k}\right)+\eta_t^{ijk}$$

$$\text{Exit: } \Pr\left[X_t^{ijk}=0|X_{t-1}^{ijk}=1\right]=c_t^{ij}+\gamma^{jk}+\boldsymbol{\lambda}'\mathbf{a}_{t-1}^{ijk}+\boldsymbol{\beta}'\left(\mathbf{z}_t^{ijk}\otimes\mathbf{mkt}^{k}\right)+\eta_t^{ijk}$$

	Low participation		High participation	
	coeff	s.e.	coeff	s.e.
	Entry			
τ_t^{jk}	-0.006	(0.002)**	0.108	(0.114)
rer_t^k	0.001	(0.001)**	0.010	(0.003)**
dem_t^k	0.005	(0.001)**	-0.002	(0.004)
Entry rate	0.007		0.017	
Share of observations	0.897		0.103	
	Exit			
τ_t^{jk}	0.179	(0.122)	-0.412	(0.350)
rer_t^k	0.005	(0.026)	0.035	(0.035)
dem_t^k	0.025	(0.027)	-0.018	(0.031)
Exit rate	0.237		0.163	
Share of observations	0.751		0.249	

Notes: Marginal effects are based on estimation of the above equation, where $\mathbf{mkt}^k$ is an indicator for whether the average participation rate across all firms and all sample years in market k is less than or greater than 10% (Participation is greater than 10% for the U.K., the U.S., and Germany). Sample is as in Table 6. Dependent variable is an indicator for entry or exit at the firm-product-market-year level. Firm-product-market fixed effects, product-market fixed effects, and baseline controls for export history are included. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

These results are reported in Table 7 which reports marginal effects by firm size, in Table 8, which reports marginal effects for low and high participation markets, and in Table 9 which reports marginal effects for different export histories. In addition to reporting the marginal effects for different groups of observations, the tables report the entry and exit rates for the relevant groups, and the share of potential entrants or incumbents they account for.

Table 9: Export entry: Marginal effects by export experience

$$\text{Entry: } \Pr\left[X_t^{ijk}=1|X_{t-1}^{ijk}=0\right]=c_t^{ij}+\gamma^{jk}+\boldsymbol{\lambda}'\mathbf{a}_{t-1}^{ijk}+\boldsymbol{\beta}'\left(\mathbf{z}_t^{ijk}\otimes\mathbf{a}_{t-1}^{ijk}\right)+\eta_t^{ijk}$$

$$\text{Exit: } \Pr\left[X_t^{ijk}=0|X_{t-1}^{ijk}=1\right]=c_t^{ij}+\gamma^{jk}+\boldsymbol{\lambda}'\mathbf{a}_{t-1}^{ijk}+\boldsymbol{\beta}'\left(\mathbf{z}_t^{ijk}\otimes\mathbf{a}_{t-1}^{ijk}\right)+\eta_t^{ijk}$$

	Entry							
	No past particip.		Last particip. at $t-2$		Last particip. at $t-3$		Last particip. bef. $t-3$	
	coeff	s.e.	coeff	s.e.	coeff	s.e.	coeff	s.e.
τ_t^{jk}	-0.005	(0.002)**	-0.343	(0.077)**	-0.145	(0.074)*	-0.056	(0.049)
rer_t^k	0.002	(0.001)**	0.004	(0.002)**	0.004	(0.002)**	0.003	(0.001)**
dem_t^k	0.005	(0.001)**	0.012	(0.002)**	0.007	(0.002)**	0.005	(0.002)**
Entry rate	0.006		0.226		0.140		0.075	
Share of observations	0.984		0.005		0.004		0.007	
	Exit							
	Tenure$_{t-1}$ = 1 yr		Tenure$_{t-1}$ = 2 yrs		Tenure$_{t-1}$ = 3 yrs		Tenure$_{t-1}$ = 4 yrs	
	coeff	s.e.	coeff	s.e.	coeff	s.e.	coeff	s.e.
τ_t^{jk}	0.058	(0.167)	-0.050	(0.173)	0.130	(0.198)	0.245	(0.249)
rer_t^k	0.014	(0.023)	0.018	(0.023)	0.012	(0.023)	0.016	(0.023)
dem_t^k	0.018	(0.025)	0.013	(0.025)	0.011	(0.025)	0.012	(0.025)
Exit rate	0.379		0.313		0.208		0.221	
Share of observations	0.205		0.123		0.079		0.051	
	Tenure$_{t-1}$ = 5 yrs		Tenure$_{t-1}$ = 6 yrs		Tenure$_{t-1}$ = 7+ yrs		Censored entry	
	coeff	s.e.	coeff	s.e.	coeff	s.e.	coeff	s.e.
τ_t^{jk}	0.154	(0.318)	-0.174	(0.346)	0.536	(0.450)	0.199	(0.137)
rer_t^k	0.018	(0.023)	0.013	(0.024)	0.014	(0.024)	0.014	(0.023)
dem_t^k	0.024	(0.025)	0.009	(0.025)	0.009	(0.025)	0.012	(0.025)
Exit rate	0.173		0.139		0.138		0.135	
Share of observations	0.037		0.026		0.035		0.444	

Notes: Marginal effects are based on estimation of the above equation, where $\mathbf{a}_{t-1}^{ijk}$ indicates the baseline controls for export history. Sample is as in Table 6. Dependent variable is an indicator for entry or exit at the firm-product-market-year level. Firm-product-market fixed effects, product-market fixed effects, and baseline controls for export history are included. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

Averaging of marginal effects across all observations does partially obscure entry responsiveness to tariffs. For large firms, and for observations with a history of recent export participation, entry is more responsive to tariff reductions than it is on average. For example, a reduction in tariffs from 10% to 0 increases the entry rate for large firms from 3.2% to 4%, and increases the entry rate for observations with the most recent history of past participation from 23% to 26%. These increases in entry rates are proportionately larger than the

increase for the sample as a whole. In contrast, entry responsiveness to real exchange rates remains modest. Meanwhile, there is no indication that export exit is sensitive to tariffs or real exchange rates in any of the subgroups we look at.

5.1.2 Entry and exit robustness

We perform a variety of additional robustness checks on our baseline analysis, varying the specification, the sample, and measurement of the independent variables. Table 10 reports the results from varying the specification in a variety of ways.

Table 10: Export entry and exit: Specification robustness

	(1)		(2)		(3)		(4)	
	No history controls		Shocks in differences		Conditional Logit		OLS with mkt f.e.	
	coeff	s.e.	coeff	s.e.	coeff	s.e.	coeff	s.e.
	Entry							
τ_t^{jk}	-0.006	(0.002)**	-0.019	(0.005)**	-2.52	(0.31)**	-0.004	(0.001)**
rer_t^k	0.002	(0.001)**	0.004	(0.001)**	0.28	(0.13)**	0.002	(0.001)
dem_t^k	0.005	(0.001)**	0.005	(0.001)**	0.44	(0.14)**	0.007	(0.001)**
History controls.	no		yes		yes		yes	
Firm-prod-yr f.e.	yes		yes		yes		yes	
Prod-mkt f.e.	yes		yes		no		no	
Mkt f.e.	no		no		yes		yes	
N	2,383,762		2,272,151		177,747		2,385,238	
R^2 or Pseudo R^2	0.29		0.29		0.18		0.25	
R^2-adjusted	0.24		0.24				0.22	
Log Pseudolikelihood					-34305.882			
Entry rate	0.008		0.008		0.108		0.008	
	Exit							
τ_t^{jk}	0.116	(0.118)	-0.469	(0.388)	1.88	(0.51)**	0.335	(0.072)**
rer_t^k	0.012	(0.023)	0.075	(0.032)**	-0.05	(0.21)	0.009	(0.022)
dem_t^k	0.001	(0.025)	-0.085	(0.079)	-0.12	(0.21)	0.017	(0.023)
History controls.	no		yes		yes		yes	
Firm-prod-yr f.e.	yes		yes		yes		yes	
Prod-mkt f.e.	yes		yes		no		no	
Mkt f.e.	no		no		yes		yes	
N	70,189		68,047		44,585		72,233	
R^2 or Pseudo R^2	0.60		0.61		0.22		0.54	
R^2-adjusted	0.48		0.50				0.45	
Log Pseudolikelihood					-11827.549			
Exit rate	0.218		0.216		0.252		0.234	

Notes: Products are defined based on the concordance of Prodcom and CN product definitions as described in Sections 2.2 and 2.3. Sample period is 1998-2009. Only potential entrants are included in the entry regression, and potential exiters in the exit regression. Dependent variable is an indicator for entry or exit at the firm-product-market-year level. Baseline controls for export history are included where noted. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

Responses to shocks are almost unaffected by dropping export history controls. Results are also similar when we include shocks in differences rather than in levels, though entry appears more sensitive to tariffs in this specification than in the baseline. This is consistent

with our prior that short-run entry (and exit) responses to a shock are bigger than long-run responses. Meanwhile, in this specification there are perverse responses of exit to real exchange rates.[22]

We also estimate a conditional logit model, where we include firm-product-year effects to control for marginal cost. For computational reasons, we cannot control for product-market effects in this specification, so we include market dummies. This implies that, in contrast to the baseline specification, coefficients on tariffs are identified from cross-sectional as well as time-series variation. We cannot calculate marginal effects in this specification, as they depend on the firm-product-year effects which are conditioned out rather than estimated. But the point estimates in the entry equation show broadly similar patterns of significance and relative magnitude to the baseline, while higher tariffs are found to increase the probability of exit. Estimation of a linear probability model with market rather than product-market fixed effects suggests that this may be partially due to the use of cross-sectional as well as time-series variation in tariffs. To sum up, results appear to be broadly robust to these variations on the specification.

Table 11 reports the results based on varying the estimation sample. We first restrict the sample to firm-product-years where the firm actually produces the relevant product.[23] We then restrict the sample to the period 2000-2009, since export participation may be censored by the relatively poorer match between Customs and CIP data before 2000. We restrict the sample to Extrastat countries, for which there is a very low threshold for reporting exports, and then to Extrastat countries plus exports to Intrastat countries by firms whose total Intrastat exports exceed the Intrastat reporting threshold. Finally, we exclude Eurozone countries from the sample. Results are broadly unchanged from the baseline in all of these variations.

Table 12 reports the results based on varying how we measure the independent variables of interest. We take liberal and strict approaches to the measurement of tariffs, and results do not change much. We also split the real exchange rate into its nominal exchange rate and price components. Entry and exit responses to these two components are very similar, despite the fact that the nominal exchange rate is much more volatile than the price component of real exchange rates.

[22]Findings are very similar when we include also lagged differences, and coefficients on lagged differences in tariffs and real exchange rates are not significantly different from zero.

[23]In the baseline sample, we require only that the firm produce the product in at least one year in the sample.

Table 11: Export entry and exit: Sample robustness

	(1)		(2)		(3)		(4)		(5)	
	Only products produced at t		Sample 2000-2009		Extrastat only		Extrastat & uncens. Intrastat		No Eurozone countries	
	coeff	s.e.	coeff	s.e.	coeff	s.e.	coeff	s.e.	coeff	s.e.
	Entry									
τ_t^{jk}	-0.009	(0.003)**	-0.006	(0.003)**	-0.009	(0.002)**	-0.007	(0.003)**	-0.008	(0.002)**
rer_t^k	0.003	(0.001)**	0.003	(0.001)**	0.001	(0.001)**	0.002	(0.001)**	0.002	(0.001)**
dem_t^k	0.005	(0.001)**	0.005	(0.001)**	0.002	(0.001)**	0.005	(0.001)**	0.004	(0.001)**
History controls	yes		yes		yes		yes		yes	
Firm-prod-yr f.e.	yes		yes		yes		yes		yes	
Prod-mkt f.e.	yes		yes		yes		yes		yes	
N	1,700,109		2,082,076		1,211,492		1,474,814		1,500,336	
R^2	0.30		0.29		0.31		0.31		0.28	
R^2-adjusted	0.25		0.24		0.23		0.24		0.21	
Entry rate	0.009		0.008		0.006		0.012		0.007	
	Exit									
τ_t^{jk}	0.133	(0119)	-0.108	(0.116)	0.062	(0.128)	0.127	(0.116)	0.067	(0.123)
rer_t^k	-0.006	(0.024)	-0.003	(0.027)	0.028	(0.033)	0.012	(0.023)	0.026	(0.029)
dem_t^k	-0.002	(0.026)	0.020	(0.035)	-0.135	(0.044)**	0.013	(0.025)	-0.062	(0.034)*
History controls	yes		yes		yes		yes		yes	
Firm-prod-yr f.e.	yes		yes		yes		yes		yes	
Prod-mkt f.e.	yes		yes		yes		yes		yes	
N	61,232		51,561		15,756		67,862		29,608	
R^2	0.61		0.64		0.60		0.60		0.61	
R^2-adjusted	0.49		0.52		0.43		0.49		0.44	
Exit rate	0.209		0.230		0.286		0.216		0.244	

Notes: Products are defined based on the concordance of Prodcom and CN product definitions as described in Sections 2.2 and 2.3. Except where noted, sample period is 1998-2009. Only potential entrants are included in the entry regression, and potential exiters in the exit regression. Dependent variable is an indicator for entry or exit at the firm-product-market-year level. Baseline controls for export history are included. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

Table 12: Export entry and exit: Robustness to measurement of independent variables

	(1)		(2)		(3)	
	Liberal tariff		Strict tariff		Split RER	
	coeff	s.e.	coeff	s.e.	coeff	s.e.
	Entry					
τ_t^{jk}	-0.006	(0.002)**	-0.007	(0.002)**	-0.006	(0.002)**
rer_t^k	0.003	(0.001)**	0.002	(0.001)**		
x_t^k					0.002	(0.001)**
cpi_t^k					0.002	(0.001)**
dem_t^k	0.005	(0.001)**	0.005	(0.001)**	0.005	(0.001)**
History controls.	yes		yes		yes	
Firm-prod-yr f.e.	yes		yes		yes	
Prod-mkt f.e.	yes		yes		yes	
N	2,383,762		2,364,153		2,383,762	
R^2	0.29		0.29		0.29	
R^2-adjusted	0.24		0.24		0.24	
In-sample entry rate	0.008		0.008		0.008	
	Exit					
τ_t^{jk}	0.068	(0.102)	0.162	(0.121)	0.126	(0.116)
rer_t^k	0.011	(0.022)	0.018	(0.023)		
x_t^k					0.017	(0.024)
cpi_t^k					0.015	(0.024)
dem_t^k	-0.004	(0.024)	0.021	(0.025)	0.012	(0.024)
History controls.	yes		yes		yes	
Firm-prod-yr f.e.	yes		yes		yes	
Prod-mkt f.e.	yes		yes		yes	
N	74,712		69,771		70,189	
R^2	0.60		0.61		0.60	
R^2-adjusted	0.48		0.49		0.49	
In-sample exit rate	0.224		0.218		0.218	

Notes: Products are defined based on the concordance of Prodcom and CN product definitions as described in Sections 2.2 and 2.3. Sample period is 1998-2009. Only potential entrants are included in the entry regression and potential exiters in the exit regression. Dependent variable is an indicator for entry or exit at the firm-product-market-year level. Baseline controls for export history are included. Liberal tariff measure includes product-market-years where there is variation in tariffs at the sub-HS6 level, using the simple average of tariffs within the HS6 as the tariff measure. Strict tariff measure excludes product-market-years where there is tariff averaging due to concordance of HS6 over time. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

5.2 Revenue

Table 13 reports the results from estimating our baseline export revenue equation, equation (5). We report only the coefficients on the shock variables. Full results, including coefficients on export history controls, are reported in the Appendix. There are two sets of estimated elasticities: those for observations with a long history of exporting (tenure of 6+ years, or entry censored by the beginning of the sample), and those for observations with a short history of exporting (1-5 years).

We focus on the elasticities for observations with long histories of exporting for which selection bias is likely to be less severe. For these observations, the elasticity of revenue with respect to tariffs is equal to -3.21, and the elasticity of revenue with respect to the real

Table 13: Export revenue responses to shocks

	coeff	s.e.
	Long export history	
τ_t^{jk}	-3.21	(0.66)**
rer_t^k	0.54	(0.09)**
dem_t^k	0.35	(0.09)**
	Short export history	
τ_t^{jk}	0.72	(0.58)
rer_t^k	0.50	(0.08)**
dem_t^k	0.26	(0.09)**
Export history controls	yes	
Firm-prod-yr f.e.	yes	
Prod-mkt f.e.	yes	
N	191,780	
R^2	0.77	
R^2-adjusted	0.68	

Notes: Products are defined based on the concordance of CN product definitions as described in Section 2.2. Sample period is 1997-2009. Dependent variable is log revenue in Euro at the firm-product-market level. Export history controls include indicators for export tenure (topcoded at 7 years), and an indicator for censored export tenure. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

exchange rate is equal to 0.54. Both elasticities are statistically different from zero. Their signs are as predicted by theory: revenue increases when tariffs in an export market fall, and when the domestic real exchange rate depreciates against a destination market. The two elasticities are statistically different from each other in absolute value. The absolute value of the elasticity of revenue with respect to tariffs is nearly six times greater than the elasticity with respect to real exchange rates, and they are statistically different at the 1% level.

In contrast, for observations with short export histories, the coefficient on tariffs is positive rather than negative, and not significantly different from zero, and the coefficient on the real exchange rate is marginally lower than in the low exit probability case. This is consistent with these observations being more subject to selection bias than the long history observations. The fact that the coefficient on tariffs is affected more than the coefficient on real exchange rates is consistent with the fact that participation is substantially more responsive to tariffs than to real exchange rates.

5.2.1 Heterogeneous elasticities

We also augment the baseline specification with a richer set of interactions on shocks. In Table 14 we report the results when we interact the shocks with indicators for firm size. In Table 15 we report the results when we interact the shocks with indicators for high and low participation markets. The coefficients on tariffs are those most affected by these additional interactions, while coefficients on real exchange rates are relatively unaffected. In particular, the absolute value of the elasticity of export revenue with respect to tariffs is greater for

large firms than for medium or small firms, and in high participation markets than in low participation markets. Since the probability of participation is greater for large firms, and in high participation markets, this is consistent with our hypothesis that selection bias is less severe for infra-marginal observations. Of course it may be that the price elasticity of demand varies across different groups. But it seems to us relatively unlikely that demand is systematically more elastic for large than small firms, and in markets with substantial penetration than in marginal markets.

Table 14: Export revenue: Interactions on shocks: Size

$$r_t^{ijk} = c_t^{ij} + \gamma^{jk} + \boldsymbol{\lambda}'\mathbf{a}_t^{ijk} + \boldsymbol{\beta}' low_t^{ijk} * \left(\mathbf{z}_t^{ik} \otimes \mathbf{size}_t^{i}\right) + \boldsymbol{\phi}' high_t^{ijk} * \left(\mathbf{z}_t^{ik} \otimes \mathbf{size}_t^{i}\right) + \varepsilon_t^{ijk}$$

	Small		Medium		Large	
	coeff	s.e.	coeff	s.e.	coeff	s.e.
	Long export history					
τ_t^{jk}	0.31	(0.97)	-2.58	(0.97)**	-5.17	(0.84)**
rer_t^k	0.53	(0.09)**	0.53	(0.09)**	0.53	(0.09)**
dem_t^k	0.30	(0.09)**	0.37	(0.09)**	0.39	(0.09)**
	Short export history					
τ_t^{jk}	2.13	(0.77)**	2.06	(0.74)**	-1.45	(0.77)*
rer_t^k	0.50	(0.09)**	0.52	(0.09)**	0.48	(0.09)**
dem_t^k	0.24	(0.09)**	0.27	(0.09)**	0.29	(0.09)**

Notes: Results based on estimation of the above equation, where $\mathbf{size}_t^i$ is an indicator for firm size. Small firms are those with fewer than 100 employees in the previous period. Medium firms had 100-249 employees in the previous period. Large firms had 250+ employees in the previous period. Products are defined based on the concordance of CN product definitions as described in Section 2.2. Sample period is 1997-2009. Dependent variable is log revenue in Euro at the firm-product-market level. Firm-product-year effects, product-market effects and baseline export history controls are included. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

Table 15: Export revenue: Interactions on shocks: low vs high participation markets

$$r_t^{ijk} = c_t^{ij} + \gamma^{jk} + \boldsymbol{\lambda}'\mathbf{a}_t^{ijk} + \boldsymbol{\beta}' low_t^{ijk} * \left(\mathbf{z}_t^{ik} \otimes \mathbf{mkt}^{k}\right) + \boldsymbol{\phi}' high_t^{ijk} * \left(\mathbf{z}_t^{ik} \otimes \mathbf{mkt}^{k}\right) + \varepsilon_t^{ijk}$$

	Low particip		High particip	
	coeff	s.e.	coeff	s.e.
	Long export history			
τ_t^{jk}	-2.43	(0.66)**	-14.18	(6.90)**
rer_t^k	0.34	(0.10)**	0.78	(0.15)**
dem_t^k	0.48	(0.10)**	-0.02	(0.13)
	Short export history			
τ_t^{jk}	0.86	(0.59)	-4.22	(7.35)
rer_t^k	0.35	(0.10)**	0.74	(0.14)**
dem_t^k	0.42	(0.10)**	-0.10	(0.13)

Notes: Results based on estimation of the above equation, where $\mathbf{mkt}^k$ is an indicator for whether the average participation rate across all firms and all sample years in market k is less than or greater than 10% (Participation is greater than 10% for the U.K., the U.S., and Germany). Products are defined based on the concordance of CN product definitions as described in Section 2.2. Sample period is 1997-2009. Dependent variable is log revenue in Euro at the firm-product-market level. Firm-product-year effects, product-market effects and baseline export history controls are included. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

5.2.2 Revenue robustness

As with entry and exit, we perform a number of robustness tests, varying the specification, the estimation sample, and how we measure the shocks. Table 16 reports the results from varying the specification. In the first column, we estimate a single set of elasticities with respect to shocks, pooling across all export observations. In this specification, we do not find any statistically significant response of exports to tariffs, while responses to real exchange rates are similar to the baseline. Results are very similar when we further drop export history controls from the regression. This is consistent with the hypothesis that selection on the transitory component of idiosyncratic demand is an issue, especially in estimating revenue responses to tariffs.

Table 16: Export revenue: Specification robustness

	(1)		(2)		(3)	
	No interactions with tenure		No history or interactions		Shocks in differences	
	coeff	s.e.	coeff	s.e.	coeff	s.e.
	All					
τ_t^{jk}	-0.44	(0.57)	-0.43	(0.57)		
rer_t^k	0.51	(0.09)**	0.56	(0.09)**		
dem_t^k	0.33	(0.09)**	0.49	(0.10)**		
	Long export history					
τ_t^{jk}					-0.66	(1.55)
rer_t^k					0.44	(0.16)**
dem_t^k					0.69	(0.22)**
	Short export history					
τ_t^{jk}					1.05	(1.53)
rer_t^k					0.19	(0.16)
dem_t^k					0.96	(0.21)**
Export history controls	yes		no		yes	
Firm-prod-yr f.e.	yes		yes		yes	
Prod-mkt f.e.	yes		yes		no	
N	191,780		191,780		124,298	
R^2	0.77		0.75		0.37	
R^2-adjusted	0.68		0.65		0.22	

Notes: Products are defined based on the concordance of CN product definitions as described in Section 2.2. Sample period is 1997-2009. Dependent variable is log revenue in Euro at the firm-product-market level. Where noted, baseline export history controls are included. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

In the third column of Table 16, we report results based on a differenced version of the baseline specification. Unlike the levels specification, we do not find any statistically significant response of revenue to tariffs, even for long export history observations. The response of revenue to real exchange rates for these observations is similar to the baseline. In interpreting this result, it is important to remember that there is relatively little year-on-year time-series variation in tariffs compared to real exchange rates, and given the very large

standard error on the coefficient on tariffs, the point estimate is not significantly different from that from our baseline fixed effects specification.[24]

Table 17 reports the results from varying the estimation sample. Restricting the sample to the period 2000-2009 for which the match between the CIP and Customs data is best does not much affect the results. Although censoring of participation in Intrastat countries is less likely to be problematic in examining export revenue than export participation, we examine what happens when we restrict the sample to Extrastat countries. Doing so restricts the sample to low participation markets, so it is perhaps not surprising that the coefficients on both tariffs and real exchange rates are closer to zero than in the baseline. We also report results restricting the sample to Extrastat countries, and to exports to Intrastat countries by firms whose total exports to Intrastat countries exceed the reporting threshold. Results are very similar to the baseline. Finally, we restrict the sample to non-Eurozone countries. Again, this restricts the sample to low participation countries, and coefficients on both tariffs and real exchange rates are closer to zero than in the baseline. Broadly speaking, however, results are similar across different samples, in the sense that in absolute value, export revenue is much more elastic with respect to tariffs than it is with respect to real exchange rates.

Table 17: Export revenue: Sample robustness

	(1)		(2)		(3)		(4)	
	Sample 2000-2009		Extrastat countries only		Extrastat & uncens. Intrastat		No Eurozone countries	
	coeff	s.e.	coeff	s.e.	coeff	s.e.	coeff	s.e.
	Long export history							
τ_t^{jk}	-3.83	(0.98)**	-2.35	(0.71)**	-3.06	(0.66)**	-2.55	(0.69)**
rer_t^k	0.59	(0.10)**	0.38	(0.13)**	0.51	(0.09)**	0.43	(0.11)**
dem_t^k	0.46	(0.12)**	0.75	(0.15)**	0.45	(0.10)**	0.75	(0.14)**
	Short export history							
τ_t^{jk}	0.68	(0.77)	0.74	(0.64)	0.77	(0.59)	0.61	(0.62)
rer_t^k	0.57	(0.10)**	0.18	(0.13)	0.46	(0.09)**	0.22	(0.11)**
dem_t^k	0.36	(0.12)**	0.58	(0.15)**	0.36	(0.09)**	0.57	(0.14)**
Export history controls	yes		yes		yes		yes	
Firm-prod-yr f.e.	yes		yes		yes		yes	
Prod-mkt f.e.	yes		yes		yes		yes	
N	151,567		132,116		160,289		84,909	
R^2	0.78		0.83		0.77		0.75	
R^2-adjusted	0.68		0.76		0.68		0.61	

Notes: Products are defined based on the concordance of CN product definitions as described in Section 2.2. Except where noted, sample period is 1997-2009. Dependent variable is log revenue in Euro at the firm-product-market level. Baseline export history controls are included. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

[24]Including lagged differences in this regression leads to an increase in standard errors on all variables.

Table 18: Export revenue: Robustness to measurement of independent variables

	(1)		(2)		(3)	
	Liberal tariff		Strict tariff		Split RER	
	coeff	s.e.	coeff	s.e.	coeff	s.e.
	Long export history					
τ_t^{jk}	-2.35	(0.56)**	-4.04	(0.71)**	-3.17	(0.66)**
rer_t^k	0.46	(0.08)**	0.53	(0.09)**		
x_t^k					0.59	(0.09)**
cpi_t^k					0.75	(0.10)**
dem_t^k	0.42	(0.09)**	0.29	(0.09)**	0.40	(0.09)**
	Short export history					
τ_t^{jk}	1.37	(0.50)**	0.21	(0.63)	0.64	(0.58)
rer_t^k	0.43	(0.08)**	0.50	(0.09)**		
x_t^k					0.56	(0.09)**
cpi_t^k					0.20	(0.11)**
dem_t^k	0.33	(0.09)**	0.20	(0.09)**	0.32	(0.09)**
Export history controls	yes		yes		yes	
Firm-prod-yr f.e.	yes		yes		yes	
Prod-mkt f.e.	yes		yes		yes	
N	208,953		190,111		191,780	
R^2	0.76		0.77		0.77	
R^2-adjusted	0.67		0.68		0.68	

Notes: Products are defined based on the concordance of CN product definitions as described in Section 2.2. Except where noted, sample period is 1997-2009. Dependent variable is log revenue in Euro at the firm-product-market level. Liberal tariff measure includes product-market-years where there is variation in tariffs at the sub-HS6 level, using the simple average of tariffs within the HS6 as the tariff measure. Strict tariff measure excludes product-market-years where there is tariff averaging due to concordance of HS6 over time. Baseline export history controls are included. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

Table 18 reports the results from varying how we measure the shock variables. Unsurprisingly, the coefficient on tariffs varies with the precision of the tariff measure we use. We also investigate whether the sensitivity of export revenue to the nominal exchange rate component of the real exchange rate differs from sensitivity to the price component. The point estimates are not significantly different from each other. We do not find evidence taht responses to the less volatile component of the real exchange rate are significantly greater than responses to the more volatile component.

To summarize, focusing on the set of observations least likely to be subject to selection bias, we find that export revenue responds to tariffs and real exchange rates in the direction predicted by economic theory. Revenue is systematically more elastic with respect to tariffs (in absolute value) than it is with respect to real exchange rates. We do observe heterogeneity in estimated elasticities across different groups of observations. However this heterogeneity is consistent with the possibility that selection bias is more severe for some groups of observations than others, and our empirical approach does not allow us to say with precision whether the underlying elasticities are different.

6 Adding up

As noted in the Introduction, estimating firm-product-market-level responses allows us to reject some potential explanations for the different responses of aggregate exports to tariffs and to real exchange rates found in the literature. But it is also of interest to know whether our micro estimates are consistent with the responses of aggregate exports to these variables. So we now use our estimates to calculate implied elasticities of aggregate exports with respect to changes in tariffs and real exchange rates.

The impact of entry on aggregate exports depends on how large is the export revenue of entrants relative to incumbents, the evolution of the export revenue of entrants post-entry, and how likely recent entrants are to exit. In addition, the impact of exit in response to shocks depends on how large is the export revenue of exiters relative to that of continuing exporters. A structural model would take account of all these factors. In the absence of a structural model, we use a regression strategy to document these features of post-entry export dynamics, and take account of them in aggregating up.

Our aggregate tariff and real exchange rate experiments are then as follows. We consider two initially identical hypothetical representative export markets, and suppose that domestic firms face a reduction in the tariff from 10% to 0 in one of the markets, but not the other, or a 10% depreciation in the home currency against one of the markets, but not the other. We calculate what happens to aggregate exports to the affected market relative to the unaffected market, *holding fixed domestic costs and the number of firms*. We calculate responses on impact, and up to six years after the shock. We also calculate long run responses. The elasticity of aggregate exports (EX) at horizon t with respect to a shock is defined as:

$$elasticity_{agg} = \frac{\ln EX_t^{change} - \ln EX^{nochange}}{\ln(1) - \ln(1.1)},$$

where *change* refers to the market where the variable of interest changes, and *nochange* refers to the market where it does not change.

6.1 Post-entry export dynamics

In estimating post-entry export dynamics, we focus on average behavior conditional on costs since our aggregate exercise holds costs fixed. We estimate average behavior over the period, controlling for shocks to tariffs and macro variables.[25]

[25] Our approach here to characterizing post-entry revenue and exit dynamics follows Fitzgerald, Haller and Yedid-Levi (2017).

We start with export revenue. We regress the log of export revenue on a set of indicator variables for tenure (topcoded at 7 years), interacting these variables with indicator variables for how long the relevant export episode lasts (also topcoded at 7 years). This allows both initial exports, and the evolution of exports post-entry, to differ between short and long export episodes. The length of an export episode acts like a proxy for the persistent component of unobserved idiosyncratic demand. We include an indicator for export episodes that are left-censored by the beginning of the sample, and for which we therefore do not observe tenure. This allows us to compare the export revenue of entrants with that of incumbents. We also include an indicator for episodes for which the completed length is censored below the top-code. Finally, we include firm-product-year effects to control for costs, and product-market-year effects to control for shocks to tariffs and macro variables.

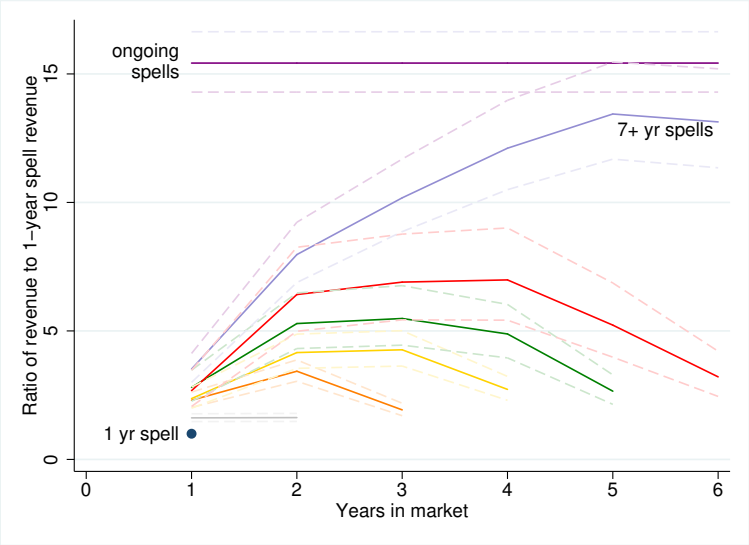
Our estimating equation is:

$$r_t^{ijk} = c_t^{ij} + \gamma_t^{jk} + \boldsymbol{\delta}' \left(\mathbf{a}_t^{ijk} \otimes \mathbf{s}_t^{ijk} \right) + \mathbf{cens}_t^{ijk} + \varepsilon_t^{ijk}, \tag{6}$$

Here, r_t^{ijk} is the log of export revenue, c_t^{ij} is a firm-product-year fixed effect, γ_t^{jk} is a product-market-year fixed effect, $\mathbf{a}_t^{ijk}$ is the vector of indicator variables for market tenure, and $\mathbf{s}_t^{ijk}$ is the vector of indicator variables for the completed length of the relevant export episode. The symbol $\otimes$ indicates the Kronecker product, but redundant interactions between $\mathbf{a}_t^{ijk}$ and $\mathbf{s}_t^{ijk}$ are dropped. The vector $\mathbf{cens}_t^{ijk}$ contains the indicators for censored entry and censored exit. The excluded category is export episodes that last only one year.

Full results from estimating equation (6) are reported in the Appendix. Figure 5 graphs the export trajectories obtained by taking the exponents of our estimates of $\boldsymbol{\delta}$. Broadly speaking, export spells that start bigger last longer. There is substantial growth of export revenue in the first five years of the longest-lasting export spells, while short-lived export spells grow more slowly, and exhibit hump-shaped dynamics, with export revenue falling prior to exit.

Figure 5: Post-entry dynamics of revenue



Notes: Figure shows evolution of revenue at the firm-product-market level with market tenure, allowing trajectories to differ based on the completed length of the relevant export episode. Revenue in export episodes lasting 1 year is normalized to 1. Ongoing spells are spells for which entry is censored by the beginning of the sample. Trajectories are estimated conditional on firm-product-year and product-market-year effects. Products are defined based on the concordance of CN product definitions as described in Section 2.2. Sample period is 1997-2009. 95% confidence intervals based on robust standard errors are plotted. Source: CSO and authors' calculations.

Next, we document the behavior of the exit hazard. We regress an indicator for exit on a vector of indicator variables for lagged export tenure (topcoded at 7) and an indicator for observations where entry is censored by the beginning of the sample. As in the case of revenue, we include firm-product-year effects to control for costs, and product-market-year effects to control for shocks to tariffs and macro variables. Our estimating equation is:

$$\Pr\left[X_t^{ijk} = 0 | X_{t-1}^{ijk} = 1\right] = c_t^{ij} + \gamma_t^{jk} + \boldsymbol{\lambda}' \mathbf{a}_{t-1}^{ijk} + cens_{t-1}^{ijk} + \varepsilon_t^{ijk}, \tag{7}$$

Here, X_t^{ijk} is the indicator for participation, c_t^{ij} is a firm-product-year fixed effect, γ_t^{jk} is a product-market-year fixed effect, $\mathbf{a}_{t-1}^{ijk}$ is a vector of indicator variables for lagged market tenure, and $cens_{t-1}^{ijk}$ is an indicator for censored entry. The excluded category is observations with lagged market tenure of 1.

The results are reported in Table 19. Exit rates decline with tenure in a market, and exit rates for observations where entry is censored are comparable to those for observations with market tenure of 6 years or more.

Table 19: Exit and past participation

Tenure	coeff.	s.e.
2 years	-0.147	(0.010)**
3 years	-0.206	(0.012)**
4 years	-0.241	(0.013)**
5 years	-0.218	(0.014)**
6 years	-0.263	(0.015)**
7+ years	-0.265	(0.013)**
censored	-0.272	(0.009)**
firm-prod-yr f.e.	yes	
prod-mkt-yr f.e.	yes	
N	47,592	
R^2	0.65	
R^2-adj	0.47	

Notes: Products are defined based on the concordance of Prodcom and CN product definitions as described in Sections 2.2 and 2.3. Sample period is 1998-2009. Only potential exiters are included in the regression. This includes exiting firms. Dependent variable is an indicator for exit at the firm-product-market level in the next year. Omitted category is observations with market tenure equal to 1 year. Censored indicates that market tenure is censored because exporting was ongoing at the start of the sample. This is extended also to observations observed to "enter" in 1997, because of the poor quality of the match between CIP and customs data in 1996. Robust standard errors calculated. ** significant at 5%, * significant at 10%. Source: CSO and authors' calculations.

6.2 Implications for aggregate export elasticities

As a first step in calculating aggregate export elasticities, we make assumptions to obtain aggregate exports to a representative market in the absence of any shock. We assume that there is a fixed number of firms in the domestic country. We assume that the rate of entry of non-participants into a representative market is equal to 0.008, which is the average entry rate for potential entrants across all the markets in our sample at the firm-product-market level (see Table 3). We assume that the rate of exit depends on market tenure. For market participants with one year of market tenure, we assume an exit rate of 45% (see Table 5). Given this 1-year exit rate, we use the coefficients on market tenure reported in Table 19 to calculate exit rates in subsequent years. Using these exit rates, we can calculate the number of export participants by export tenure and predicted export spell length.[26] We then assign an average level of exports to each tenure-spell-length cell based on the estimated initial levels and growth trajectories in Figure 5 (see Appendix for the corresponding table). Multiplying the number of participants per cell by the average exports per cell gives us total exports to the representative market in the absence of any shocks.

Next, we trace through the impact of a reduction in the tariff faced by domestic firms in the representative market from 10% to 0, or a 10% depreciation of the real exchange

[26]The participation rate we obtain is very close to the average participation rate from Table 5.

rate between the home country and that of the representative market. We hold constant the number of domestic firms as well as marginal cost for all firms. The initial conditions are given by the distribution of participants and the level of exports across tenure-spell-length cells just described. We then apply our baseline estimates of entry, exit, and revenue responses from Section 5.[27]

For entry rates, things are straightforward. The coefficients on both tariffs and real exchange rates in the baseline entry equation are statistically different from zero (see the first column of Table 6), so we apply these coefficients to the entry rate. A reduction in the tariff from 10% to 0 increases the entry rate from 0.008 to 0.0087, while a 10% depreciation increases the entry rate from 0.008 to 0.0082. This affects the formation of new cohorts.

For exit, the coefficients on tariffs and real exchange rates in our baseline exit equation are not significantly different from zero (see the second column of Table 6). This raises the question of whether we should allow for any impact of changes in these variables on exit rates. We calculate aggregate elasticities two ways: first, using the point estimates from our baseline exit equation, and second, setting exit responses to zero. In the first case, we assume that a reduction in the tariff from 10% to 0 reduces exit rates at all levels of market tenure by 0.0127 (e.g. from 0.45 to 0.4373 for spells with tenure of one year) while a 10% depreciation *increases* all exit rates by 0.0015 (e.g. from 0.45 to 0.4515 for spells with tenure of one year). Note that we apply the same change in exit rates to all incumbents, irrespective of tenure.[28] When an export spell ends up being longer or shorter under the new tariff or real exchange rate than it would have been in the absence of the shock, we assume that the revenue trajectory it follows (exclusive of the revenue response to the shock) is governed by its predicted spell length on entry.

For export revenue, we have two sets of estimated elasticities, those for observations with short export histories, and those for observations with long export histories (see Table 13). Although the estimated elasticities for observations with short histories are likely subject to selection bias, they may still allow us to capture an important feature of responses to shocks: marginal participants induced to export by a shock are likely to have lower average exports than infra-marginal participants. A crude way to capture this effect is to apply the short history elasticities to tenure-spell-length cells with short tenure (1-5 years), and the long history elasticities to cells with long tenure (6+ years). Our estimates imply that in short tenure cells average export revenue decreases in response to tariff reductions, with

[27] As noted in Section 5, short run entry and exit responses may differ from long run responses. In the Appendix we report results incorporating this.

[28] In the Appendix we report results allowing for variation in the impact by tenure.

an elasticity of 0.73, while in long tenure cells average revenue increases with an elasticity of 3.21. Meanwhile in short tenure cells, export revenue increases in response to a real depreciation, with an elasticity of 0.50, while the elasticity for long tenure cells is 0.54. We also calculate aggregate elasticities applying the long history elasticities to all cells.

Table 20 reports the results of these exercises. The top panels report responses to the tariff shock, while the bottom panels report responses to the real exchange rate shock. The left panels report responses making use of both short and long history revenue elasticities, while the right panels report responses using only long history revenue elasticities. Within each panel, the first column reports the elasticity of aggregate exports setting all responses to their baseline point estimates. The second column reports the elasticity setting exit responses to zero. The final two columns in each panel report the aggregate elasticities allowing only revenue responses, and only entry and exit responses respectively.

Table 20: Elasticities of aggregate exports with respect to shocks

	Tariff shock							
	Short and long tenure revenue elasticities				Long tenure revenue elasticities only			
	Margins							
Horizon	entry, exit, rev	entry, rev	rev	entry, exit	entry, exit, rev	entry, rev	rev	entry, exit
1	-1.68	-1.50	-1.44	-0.25	-3.45	-3.28	-3.21	-0.25
2	-1.93	-1.58	-1.44	-0.52	-3.73	-3.38	-3.21	-0.52
3	-2.15	-1.65	-1.44	-0.76	-3.97	-3.47	-3.21	-0.76
4	-2.36	-1.73	-1.44	-0.98	-4.19	-3.56	-3.21	-0.98
5	-2.55	-1.80	-1.44	-1.17	-4.38	-3.64	-3.21	-1.17
6	-2.74	-1.89	-1.44	-1.32	-4.53	-3.71	-3.21	-1.32
steady state	-3.55	-2.30	-1.44	-2.01	-5.23	-4.07	-3.21	-2.01
	Real exchange rate shock							
	Short and long tenure revenue elasticities				Long tenure revenue elasticities only			
	Margins							
Horizon	entry, exit, rev	entry, rev	rev	entry, exit	entry, exit, rev	entry, rev	rev	entry, exit
1	0.52	0.54	0.52	-0.00	0.54	0.56	0.54	-0.00
2	0.53	0.57	0.52	0.01	0.55	0.59	0.54	0.01
3	0.53	0.59	0.52	0.01	0.55	0.61	0.54	0.01
4	0.54	0.62	0.52	0.03	0.57	0.64	0.54	0.03
5	0.56	0.65	0.52	0.04	0.58	0.67	0.54	0.04
6	0.57	0.67	0.52	0.05	0.59	0.69	0.54	0.05
steady state	0.64	0.77	0.52	0.12	0.66	0.79	0.54	0.12

Notes: First column of each panel reports elasticities allowing entry, exit, and revenue to respond. Second column of each panel reports elasticities allowing entry and revenue to respond. Third column reports elasticities allowing only revenue to respond. Fourth column reports elasticities allowing only entry and exit to respond. For details of construction of aggregate elasticities, see text.

The elasticities reported in Table 20 are broadly in line with those estimated using aggregate data (see Ruhl (2008)). On impact, aggregate behavior is governed mainly by the response of export revenue conditional on participation. Over time, the contribution of participation responses grows, so long run responses do indeed differ from short run responses. But the elasticity of aggregate exports with respect to tariffs is *always* greater than the

elasticity with respect to real exchange rates - between 2.5 and 8 times as large, depending on the time horizon and the method of calculation.

7 Conclusion

The main contribution of this paper is to use production and export micro data for Ireland to estimate the responses of export entry, exit, and the export revenue of incumbent exporters to tariffs and real exchange rates. We find that export entry is three times more responsive to tariffs than to real exchange rates, while the export revenue of long-time incumbents is nearly six times more responsive to tariffs than it is to real exchange rates. Taking account of post-entry export dynamics, we find that long run responses of aggregate exports to both shocks differ from short run responses. But at all horizons, aggregate exports are at least 2.5 times more responsive to tariffs than they are to real exchange rates.

Our results conclusively reject two proposed explanations for the International Elasticity Puzzle: aggregation bias, and naive comparisons of long run elaticities (tariffs) with short run elasticities (real exchange rates). On a more positive note, recent research suggests at least two factors which may contribute to the greater responsiveness of exports to tariffs than to real exchange rates at the firm-product-market level. Our findings are consistent with these explanations. Ruhl (2008) points to the different statistical processes followed by tariffs (persistent, not very volatile) and real exchange rates (less persistent, more volatile). He points out that irreversibilities in export decisions such as sunk costs of export participation may interact with these different statistical properties to generate different participation responses to the two shocks. Meanwhile Fitzgerald, Yedid-Levi and Haller (2017) note that when expenditures on marketing and advertising take place in the destination market as in Arkolakis (2010) and Drozd and Nosal (2012), export responses to real exchange rates may be muted. This mechanism is unlikely to affect responses to tariffs, and can therefore rationalize different revenue responses to the two shocks.

Our empirical findings have important implications for a number of policy questions. They suggest only a very modest impact of "currency manipulation" on international trade. In addition, they imply that the real exchange rate adjustment necessary to offset a border adjusted corporation tax may be quite large. More generally, our findings suggest caution in assuming that combinations of alternative policies that affect firms through real exchange rates have an impact equivalent to changes in trade policy, as firms seem to respond very differently to comparable innovations in these two variables.

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