Trade and Investment under Policy Uncertainty: Theory and Firm Evidence

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ABSTRACT: Using a dynamic, heterogeneous firms model with sunk costs of exporting we show that: (i) investment and entry into export markets is reduced when trade policy is uncertain and (ii) credible preferential trade agreements (PTAs) increase trade even if applied trade barriers are currently low. We structurally estimate the effect of policy uncertainty on firm entry following Portugal's accession to the European Community in 1986 and find that (i) the trade policy reform accounted for a large fraction of the observed Portuguese exporting firms' entry and sales (ii) the accession removed uncertainty about future EC trade policies and (iii) this uncertainty channel accounted for a large fraction of the predicted growth. These results have broader implications for other PTAs and our approach can be applied to analyze other sources of policy uncertainty.

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Firms face considerable uncertainty about future conditions, which can arise from purely economic shocks—e.g. to productivity or tastes—or from policy shocks—e.g. monetary or fiscal reforms. The role of future conditions is particularly important when firms must decide on costly irreversible investments. We examine the impact of policy uncertainty on a firm's decision to invest and export to new markets, which is an interesting setting for several reasons. First, global integration has considerably increased firms' exposure to foreign policy uncertainty. Second, while most trade analysis assumes policy is deterministic, we argue that it can be quite uncertain. Trade policy shocks are not frequent but when they happen they can be large and persistent, as witnessed in the 1930's trade war.[1] Third, there is growing evidence that firms must incur sunk costs to start exporting (cf. Roberts and Tybout, 1997). To capture these elements we develop a tractable dynamic heterogeneous firms' model under policy uncertainty. We derive, estimate and quantify the impacts of current and future trade policy on investment and export decisions by combining novel firm-level and trade policy data in the context of preferential trade agreements.

A broader motive to focus on trade is the rich policy and firm data available, which has the potential to inform us about the firm impacts of other types of policy uncertainty. The basic theoretical impacts of uncertainty on investment are understood (cf. Bernanke, 1983), and recent evidence links aggregate uncertainty shocks to investment.[2] However, there is scant empirical evidence of the importance of policy uncertainty for firms, even though thousands of firms worldwide rank it as 'one of the most important constraints in doing business' (World Bank Development Report, 2005). The scant evidence is likely due to the inherent conceptual and empirical challenges surrounding this question. These include the difficulty in measuring policy uncertainty and clearly identifying its causal impact on specific investment decisions.[3] The international trade setting can help address these issues. First, it allows us to construct detailed measures of policy uncertainty that are easy-to-interpret and vary across countries, products and time. Second, we can trace the effects of these measures to specific firm investment and sales decisions that also vary along all these dimensions and can thus help address endogeneity concerns.

We build on the option value insights in Dixit (1989) and Rodrik (1991) and extend them to a context with trade policy uncertainty (TPU) and heterogenous firms. We characterize policy by using a stochastic process that generates closed form solutions and clearly guides the empirical estimation and quantification. The theoretical framework can be applied to alternative sources of TPU including the timing and credibility of large trade policy reforms, be they unilateral (e.g. China, India and Brazil's liberalizations), multilateral (e.g. completion or implementation of trade

[1] Fears of a similar war during the Great Recession of 2008-09 prompted leaders of the G-20 to repeatedly pledge that "We will not repeat the historic mistakes of protectionism of previous eras." http://www.londonsummit.gov.uk/en/summit-aims/summit-communique/ While a trade war did not break out, the initial fears were not unfounded since non-tariff barriers have often been used during downturns (Bown and Crowley, 2013). Other sources of potentially large and permanent changes in protection include (i) import bans due to safety concerns; (ii) threats of tariffs to counter exchange rate manipulation or carbon emissions in production.

[2] Bloom et al. (2007) and Bloom (2009) provide evidence that shocks to stock market volatility delay firm-level investment and attenuate its response to demand shocks.

[3] As Rodrik (1991) notes "the idea that policy instability can be detrimental to private investment is easy to accept (...) However, it is hard to deploy serious econometrics in support of the proposition".

rounds) or bilateral (e.g. passage of NAFTA).[4]

In order to identify the effect of *changes* in policy uncertainty we focus on whether and how trade agreements reduce TPU and thus affect export investments. Our empirical analysis focuses on preferential trade agreements (PTAs)—the most important source of trade policy reform in the last 20 years.[5] Importantly, many such agreements claim to 'reduce distortions to trade' *and* 'ensure a predictable environment for business planning and investment'.[6] As we will show, a key determinant for whether PTAs spur export investment is whether the tariff reductions they generate are credible or face some probability of reversal. This concern is pervasive for the many developing countries that obtain unilateral preferences from the U.S. or Europe.[7]

The empirical application focuses on Portugal's accession to the European Community (EC). This is a good setting to study the trade and investment effects of TPU for several reasons. First, the focus on a specific country and policy event allows us to cleanly identify the effect and carefully control for a number of factors. Second, the effects of TPU should be important for small, developing, open economies where trade is central both to consumers and firms.[8] So, this episode can be relevant for many developing countries currently seeking secure access to the U.S. and European markets.[9] Third, Portuguese trade increased dramatically after 1986 and that increase was largest towards the EC partners, *suggesting* that it was caused by the accession. Finally, the export expansion upon accession was characterized by considerable entry of Portuguese firms exporting into EC markets even in industries where applied tariffs did not change and we find this cannot be explained by standard determinants such as income or exchange rates, which indicates a potential role for the agreement in reducing TPU.

Our model shows how to measure, test and quantify the impact of TPU on firm entry and trade. We estimate a structural equation where export entry depends on current policy and a measure of TPU: the percent loss in profits due to a negative tariff shock that reverts tariffs to their non-preferential levels, which we observe. We find supporting evidence for several of the model's predictions by exploring variation in firm entry and policy across different export markets (EC-10 and Spain), industries and time. The estimates indicate that Portuguese exporters believed that the probability of losing preferences was zero after EC accession but positive before. So the agreement eliminated that source of TPU.

[4] U.S. trade agreements are subject to Congressional modification if the president does not have "fast-track" authority (Conconi et al, 2012), which can generate TPU.

[5] In 2014, there were 377 PTAs in force and 583 notified to the WTO. www.wto.org

[6] For examples see the Global PTA Database at <wits.worldbank.org/GPTAD>. The WTO also aims to generate "Predictability through bindings and transparency [to] promote investment (...)" www.wto.org.

[7] U.S. preferences to several developing countries are often conditional on a variety of criteria that can and have triggered non-renewal for periods up to fourteen months.

[8] A large fraction of Portuguese firms are engaged in some form of international trade—about 24%—and account for a large fraction of private sector non-agricultural employment—58% or 46% if we focus only on exporters. These figures for 1987 are based on merged information Quadros de Pessoal and Portuguese Trade statistics.

[9] For example, Peru and Colombia received U.S. preferences subject to periodic renewal and subsequently agreed to lower barriers on U.S. products to permanently secure those preferences and spur investment (cf. USITC, 2008). Similarly, several countries that seek (or sought) PTAs with the EC previously received some form of preference.

Using the estimated structural parameters we then quantify the impact of alternative policies. The trade policy changes due to accession accounted for 61% of the growth in entry and 87% of the growth in export value. If the accession had only reduced applied tariffs, but not TPU, it would have achieved only 20% of the total predicted firm entry and less than 30% for total exports. So, a substantial fraction of the growth was generated by the elimination of TPU. Moreover, about 65% of the growth in entry is due to securing previous tariff reductions, and more than half of this is accounted for by a mean preserving reduction in tariff risk.

Our work is related to research in different fields. We build on Dixit (1989) who shows that price uncertainty creates an option value of waiting before making sunk cost entry investments. These insights have mainly been used to analyze the impact of exchange rate uncertainty on (i) exports (cf. Baldwin and Krugman, 1989, and Das et al, 2007); (ii) FDI (Russ, 2007) or both (Garetto and Fillat, 2014).[10] We too develop a dynamic model of trade with sunk costs, but we focus on TPU and provide closed form solutions that guide the estimation.

The impact of trade and tax policy uncertainty when there are sunk costs of investment, has received far less attention. For example, Rodrik (1991) develops a model of capital investment when firms believe an investment tax credit may be reversed in the future. If the probability or cost of a policy reversal is high, a reform to promote investment may produce the opposite outcome. The scant empirical evidence has mostly focused on aggregate outcomes; Baker et al. (2012) find that increases in a news-based index of policy uncertainty are negatively correlated with aggregate employment, output, and investment in the U.S. Our approach is different and our trade focus aims to identify the effects of policy uncertainty shocks as cleanly as possible.

We also contribute to the debate regarding the value of trade agreements. Early empirical work on the trade impacts of PTAs delivered mixed results, e.g. Frankel (1997) reports small and sometimes negative effects of EC membership on trade between members in the 1960s and 1970s but positive ones in the 1980s and 1990s. Our approach highlights that whether PTAs increase bilateral trade depends on whether their policies are credible.[11] The role of TPU is also explored by Handley (2014), who provides evidence that WTO tariff bindings reduce TPU and thus promote export entry.[12] Limão and Maggi (forthcoming) theoretically examine the value of endogenous trade agreements under uncertainty in a perfectly competitive setting with risk averse individuals.

In section 2 we derive the structural relationship between TPU and firm investment decisions. In section 3 we describe the empirical approach, provide estimates and quantify the policy impacts. In the final section we summarize our results and discuss implications and possible extensions.

[10]There is stronger evidence for a negative effect of exchange rate volatility on FDI (cf. Campa, 1993, and Russ, 2012) than exports (cf. Das et al, 2007, and Alessandria and Choi, 2007). Impulliti et al (2013) show that firm productivity uncertainty can generate hysteresis in firms' export decisions.

[11]This can help explain why even though PTAs increase trade substantially on average (cf. Baier and Bergstrand, 2007) not all PTAs do so (Baier et al., 2007). Egger et al (2011) address the endogeneity of PTAs and zero bilateral flows in a unified way. Ruhl (2004) argues that PTAs can generate export entry by permanently lowering trade frictions, which strengthens response to future macro variable shocks.

[12]Sala et al. (2010) study the impact of WTO bindings on exports theoretically.

2 Theory

To determine the impact of TPU on export entry investments we first determine operating profits for each firm conditional on exporting. Second, we examine a firm's decision to invest to enter that market and how it is affected by policy uncertainty.

2.1 Demand and Operating Profits

The operating profit for a firm that exports a differentiated good, v, to country i is determined as follows. At the start of each period a firm can observe the policy and demand that it will face in each market before it produces, which allows us to focus on the export entry investment as the sole decision made under uncertainty. This timing, and the absence of any adjustment costs, implies that, after entry, firms simply maximize operating profits period by period and those profits can be derived similarly to standard deterministic monopolistic competition models.

One industry produces a homogeneous, freely traded good—the numeraire—and each of the remaining V industries produce differentiated goods. Each industry V has a fixed exogenous fraction μ_V of each country's total expenditure on goods and the rest is spent on the numeraire. Consumers have constant elasticity of substitution preferences over goods in each V with $\sigma = 1/(1-\rho) > 1$. Therefore a firm v faces the standard optimal demand, $q_{iv} = \tilde{A}_{iV} p_{iv}^{-\sigma}$, where $\tilde{A}_{iV}$ is an industry demand parameter, which is exogenous from the perspective of any given firm.[13] The consumer price, p_{iv}, includes trade costs. We focus on ad valorem import tariffs and note that they are product or industry specific. We denote the tariff factor that i sets on V by $\tau_{iV} \geq 1$, so free trade is represented by $\tau_{iV} = 1$. Therefore, producers of any $v \in V$ exporting to i receive p_{iv}/τ_{iV}.

The supply side is standard. There is a single factor, labor, which has constant marginal productivity in the numeraire sector, so the wage is normalized to unity. Differentiated goods are produced with a constant marginal cost, characterized by a labor coefficient of c_v, which is heterogenous across firms. The firm chooses prices to maximize operating profits in each period, $(p_{iv}/\tau_{iV} - c_v)\, q_{iv}$, leading to the standard mark-up rule over cost, $p_v = c_v/\rho$. Consumers in i face this price augmented by the tariff, that is $p_{iv} = (c_v/\rho)\,\tau_{iV}$. Using the optimal price and demand we obtain, respectively, the export revenue received by the firm and the associated operating profit:

$$p_{iv} q_{iv}/\tau_{iV} = (\tau_{iV})^{-\sigma}\, \tilde{A}_{iV}\, (c_v/\rho)^{1-\sigma} \tag{1}$$

$$\pi(c_v, \tau_{iV}) = (\tau_{iV})^{-\sigma}\, A_{iV} c_v^{1-\sigma} \tag{2}$$

where $A_{iV} \equiv \tilde{A}_{iV}\,(1-\rho)\,\rho^{\sigma-1}$ summarizes industry demand conditions in the foreign market.

[13] Under this structure $\tilde{A}_{iV} = \mu_V Y_i\,(P_{iV})^{\sigma-1}$ where P_{iV} is the standard CES price aggregator over varieties in each V sold in i and Y_i is the aggregate expenditure on goods. The standard assumption is that each firm is sufficiently small relative to the total number in industry V and country i to take into account any effect that they may have on the price index or aggregate goods' expenditure.

2.2 Firm Value, Investment and Export Entry Setup

In each industry V there is an exogenous constant mass of firms that produce for their home market and an endogenous subset of them that exports.[14] Firms face no uncertainty about their productivity when deciding whether to export (they already produce domestically). To focus on foreign demand uncertainty due to policy we assume the following. First, the single underlying source of uncertainty is the future tariff policy in the foreign market. Second, the exporting country is sufficiently small that changes in the policy it faces when selling to i have no impact on the foreign aggregate demand term, A_{iV}.[15] Given these assumptions and the constant mass of domestic producers in i, we will treat A_{iV} as constant in the theory when analyzing firms' decisions to export to this market. We can allow for unanticipated shocks to A_{iV} and will control for them empirically.

Let us first consider a benchmark case where the policy is deterministic. A firm enters a new export market if the present discounted value of its profit exceeds the sunk investment cost of entry, i.e. $\pi(c_v, \tau_{iV}) \geq K_{iV}(1-\beta)$. The discount factor $1-\beta$ reflects the exogenous probability of a death shock to exporting.[16] Each industry has a continuum of firms with heterogenous marginal cost drawn from a distribution function $G_V(c)$. Thus we can use the condition above to determine the marginal entrant without policy uncertainty.

$$c^D_{iV} = \left[\frac{(\tau_{iV})^{-\sigma}}{1-\beta} \frac{A_{iV}}{K_{iV}} \right]^{1/(\sigma-1)} \tag{3}$$

The cutoff is common to all firms that face a similar tariff and sunk cost, so for $v \in V$ all firms with $c_v < c^D_{iV}$ enter. Reductions in tariffs increase demand and thus induce entry.

Trade policy is rarely deterministic and potential exporters can choose when to invest so TPU can generate an option value of waiting. The analysis below applies for each firm v considering the decision to export to a given market i so we drop these subscripts for simplicity. We model the entry decision as an optimal stopping problem. Firms can be divided into exporters and non-exporters. The expected value of being an exporter is denoted by Π_e; such a firm exits only when hit by a "death" shock since it has no other fixed costs after it enters.[17] Non-exporters decide to enter only when the value of exporting net of the sunk entry costs exceeds the expected value of waiting, Π_w.

[14] To focus on firms' foreign market entry decisions we assume there are no fixed costs to enter the domestic market. This simplification does not affect our basic empirical results since our identification approach controls for industry-time effects and thus accounts for domestic entry into any particular industry.

[15] This assumption is reasonable in the setting we examine empirically—a small country (Portugal) exporting to the EC—but should be relaxed in other settings. In Handley and Limão (2013) we extend this model and allow for general equilibrium effects of the policy, via impacts on the price index, in order to analyze the impact of China's entry into the WTO on its exports to the US. Doing so introduces a price index adjustment, which attenuates the direct effects of tariff policy on entry decisions but does not overturn them.

[16] Similar qualitative results hold if we allow firms to place less weight on future profits due to an exogenous pure time preference factor.

[17] While the assumption of no per period fixed costs of exporting may seem extreme, Das et al. (2007) find these per period fixed costs are negligible, on average, across all sectors analyzed in their structural model of Colombian exporters. When we incorporate endogenous exit the analysis is more complex but does not change the central results nor does it add much additional insight for the empirical application we pursue.

The investment and entry decision rule for a firm with cost c can be defined as a function of the threshold tariff $\bar{\tau}$ that makes it indifferent between entering today or waiting.

$$\Pi_e(c, \bar{\tau}) - K = \Pi_w(c, \bar{\tau}) \tag{4}$$

So, any tariff below $\bar{\tau}(c)$ triggers entry by firms with cost c or lower. To characterize the impact of uncertainty on this entry cutoff we describe the policy process and derive these value functions.

2.3 Trade Policy Regimes

We model trade policy as an exogenous stochastic process that is tractable and rich enough to span alternative trade policy regimes. Each period there is some probability γ of a policy shock (e.g. a new agreement, a new government, a macroeconomic shock, etc.). If the shock occurs then a policy maker reconsiders the current policy and sets a new one denoted by τ', otherwise the policy is unchanged. Firms form expectations over future policies based on their belief of γ and a probability measure of tariff outcomes, $H(\tau')$, with support $\tau' \in [\tau^L, \tau^H]$, where τ^H is the worst case scenario. We assume that both γ and H are similar across firms within a given industry V so that entry decisions depend only on a firms' productivity relative to the industry cutoff.

The stochastic path of the policy is characterized by the current tariff, τ_t, and the "policy regime" described by the pair $\{\gamma, H\}$. Firms believe that the *regime* is time-invariant and take it as given. Our results will then apply to any given exogenous γ and H and will allow us to compare equilibrium firm behavior across different regimes. This characterization encompasses alternative regimes. When $\gamma \in (0, 1)$ we have a setting with imperfectly anticipated shocks of uncertain magnitude. Alternatively, if $\gamma = 1$ and H is degenerate at some τ' then a reform is perfectly anticipated and the government credibly committed to τ' in the following period. When $\gamma = 0$ the government has committed to the current tariff, τ_t. The setup can also be used to capture staged tariff reductions—typical in trade agreements—provided that their implementation is uncertain.

2.4 Value of Credible vs. Non-credible Policies

We now derive the entry decisions and the impact of policy changes within a credible regime ($\gamma = 0$) vs. a non-credible one ($\gamma > 0$, so a policy change is expected).

The expected value of a firm that exports at time t conditional on it observing τ_t is given by

$$\Pi_e(c, \tau_t) = \pi(c, \tau_t) + \beta[\underbrace{(1-\gamma)\Pi_e(c, \tau_t)}_{\text{No Shock}} + \underbrace{\gamma \mathbb{E}\Pi_e(c, \tau')}_{\text{Shock}}], \tag{5}$$

which includes current operating profits in the *export market* and the discounted future value.[18]

[18]We do not explicitly include profits from domestic sales because in our setting they are independent of export conditions and constant over time. We can include them by adding the same constant term to the value of exporting

With probability $1-\gamma$, there is no policy shock and the expected firm value is unchanged. With probability γ, there is a policy shock. So the third term on the RHS of (5) is the *ex-ante* expected value of exporting following a shock and $\mathbb{E}$ denotes the expectation over the H distribution.

The expected value of a firm waiting to export is given by

$$\Pi_w(c) = 0 + \beta[\underbrace{(1-\gamma)\Pi_w(c)}_{\text{No Shock}} + \underbrace{\gamma(1-H(\bar{\tau}))\Pi_w(c)}_{\text{Shock Above Trigger}} + \underbrace{\gamma H(\bar{\tau})(\mathbb{E}[\Pi_e(c,\tau') \mid \tau' \leq \bar{\tau}] - K)}_{\text{Shock Below Trigger}}]. \tag{6}$$

A non-exporter receives zero profits from that activity today; the continuation value stays at Π_w if there is no shock or the shock entails a tariff above the trigger. If a policy shock arrives, it will be below $\bar{\tau}$ with probability $H(\bar{\tau})$ and the firm will choose to pay K and start exporting.

To obtain the equilibrium values of exporting, $\Pi_e(c,\tau_t)$, and waiting, $\Pi_w(\tau_t)$, we solve the linear system given by (5), (6), $\mathbb{E}\Pi_e(c,\tau')$ and $\mathbb{E}[\Pi_e(c,\tau') \mid \tau' \leq \bar{\tau}]$ (given in Appendix A) and obtain

$$\Pi_e(c,\tau_t) = \frac{\pi(c,\tau_t)}{1-\beta(1-\gamma)} + \frac{\beta\gamma}{1-\beta}\frac{\mathbb{E}\pi(c,\tau')}{1-\beta(1-\gamma)} \tag{7}$$

$$\Pi_w(c) = \frac{\beta\gamma H(\bar{\tau}(c))}{1-\beta\left(1-\gamma H\bar{\tau}(c)\right)}\left\{\frac{\mathbb{E}[\pi(c,\tau') \mid \tau' \leq \bar{\tau}(c)]}{1-\beta(1-\gamma)} + \frac{\beta\gamma}{1-\beta}\frac{\mathbb{E}\pi(c,\tau')}{1-\beta(1-\gamma)} - K\right\} \quad \text{if } \tau_t \geq \bar{\tau}(c). \tag{8}$$

The interpretation of (7) is straightforward: after investment, the firm value from exporting conditional on τ_t equals the discounted value of expected profits. If γ were zero this would be the deterministic value. But the policy changes with probability γ and the ensuing per period expected profits are $\mathbb{E}\pi(c,\tau')$. If the current tariff is above a given firm's trigger, $\tau_t > \bar{\tau}(c)$, then it does not export today and its value, $\Pi_w(c)$, would be zero if the tariff remained above that trigger, but with probability $\gamma H(\bar{\tau})$ the tariff will fall below the trigger leading the firm to incur K and export. The expected value of exporting is then captured by the remaining terms in brackets.

Using (7) we see that reductions in the current tariff, τ_t, increase the firm value, Π_e, at any γ. Moreover, that impact is higher under a credible policy regime ($\gamma = 0$) then when the policy is expected to change ($\gamma > 0$). This complementarity between reductions in current tariffs and uncertainty is one possible reason why countries that obtain temporary preferential tariffs spend considerable resources trying to eliminate any uncertainty about their future reversal.

2.5 Policy Impacts on Investment and Entry

We now derive the impact of different policy changes on entry. Using (7), (8), and (4) we determine the threshold tariff $\bar{\tau}(c)$ such that a firm with cost c is indifferent between waiting or exporting. Assuming $G(c)$ is strictly increasing we can invert the threshold function $\bar{\tau}\left(c_t^U\right) = \tau_t$ to uniquely determine the cutoff cost for any given current tariff. In appendix A we show that c_t^U equals the

and to the value of waiting, which will change none of the subsequent results.

product of an uncertainty factor and the deterministic cutoff

$$c_t^U = U_t \times c_t^D \tag{9}$$

$$U_t \equiv \left[\frac{1-\beta\left(1-\gamma\omega\left(\tau_t\right)\right)}{1-\beta\left(1-\gamma\right)}\right]^{\frac{1}{\sigma-1}} \leq 1 \tag{10}$$

We highlight two properties of the uncertainty factor U_t, which captures the frequency of policy shock arrivals and expectations about future tariffs. First, if $\gamma = 0$ then $U_t = 1$ and we are back to the deterministic cutoff, c_t^D. Second, if $\gamma > 0$ then there is lower entry whenever the current tariff is below the maximum, i.e. $\tau_t < \tau^{\max}$. To see this note that when $\gamma > 0$, we obtain $c_t^U < c_t^D$ if and only if $\omega\left(\tau_t\right) < 1$. In appendix A we show that $\omega\left(\tau_t\right) - 1$ is the expected growth in operating profits conditional on a shock that increases tariffs, which is negative iff $\tau_t < \tau^{\max}$. This expected profit loss term will be central in the empirical application so we record it here:

$$\omega\left(\tau_t\right) - 1 = -\left(1 - H(\tau_t)\right)\frac{\tau_t^{-\sigma} - \mathbb{E}(\tau^{-\sigma} \mid \tau \geq \tau_t)}{\tau_t^{-\sigma}} \leq 0. \tag{11}$$

We make three observations First, because this is a proportional loss of operating profits it is independent of a firm's productivity. Second, when tariffs are the only source of uncertainty, we can use information on the current tariff and a measure of the expectation of future tariffs *above* τ_t to construct ω. So, even though the policy shock can trigger a lower or higher tariff, it is only the possibility of higher tariffs that affects entry.[19] Third, this term varies with tariffs across different industries, which will prove useful to identify the impact of uncertainty.

We conclude this section by noting that for any $\tau_t < \tau^{\max}$ entry is monotonically decreasing over all values of γ. This has two implications for the empirical analysis. First, we can test if any given event, e.g. an agreement, impacts entry via uncertainty without assumptions on whether γ was zero before or after that event.[20] Second, if the current tariff is at the long-run mean then reductions in γ can be interpreted as mean risk compressions of the tariff (see appendix A). This insight will allow us to decompose the quantitative effects of moving to a more credible policy regime into two components: pure risk and long-run mean changes of the policy.[21]

3 Evidence

We explore the theoretical framework to address two questions. What are the effects of current policy and uncertainty on export entry? Do trade agreements reduce uncertainty?

[19] This is an example of the "bad news" principle first identified by Bernanke (1983).

[20] More specifically, $d\ln c_t^U/d\gamma < 0$ if $\tau_t < \tau^{\max}$ as derived in Appendix A.

[21] In supplementary appendix D.1 we also show that for *any* τ_t above the minimum there exist mean preserving compressions in the H distribution that increase entry.

3.1 Policy and Institutional Setting

We focus on Portugal's accession to the European Community (EC) in 1986. One reason for doing so is that this agreement reduced some tariffs that Portuguese exports faced *and* secured pre-existing preferences to the EC markets and Spain, which joined the EC in 1986.

Portugal's EC accession was signed on June 1985 and implemented on March 1^{st}, 1986. EC tariffs on Portuguese industrial goods did *not* change in 1986, they remained duty-free, as they had been since 1977. So the trade policy impact of the 1986 accession on such products will *not* be due to changes in applied tariffs. We will test if the impact of accession on those products was to make those pre-existing preferences more credible, i.e. whether the accession lowered the probability that Portugal lost its EC preferences and faced the 8% average tariffs that other countries—those without preferential access to the EC—faced.[22]

Spain also joined the EC in 1986 and was required to liberalize its industrial tariffs against Portugal. Spanish tariffs on Portuguese goods prior to the agreement were 7.9%. So the impact of accession on Portuguese exports towards Spain will reflect a reduction in applied tariffs, but it will also reflect an uncertainty reduction component related to securing pre-existing preferences. In 1980 Spain signed an agreement that began a partial liberalization of its tariffs towards certain countries including Portugal. But this agreement contained no definite timetable or scheduled reductions and was potentially incompatible with GATT's requirement that PTAs implement zero tariffs on substantially all trade.[23] Thus, before 1986, Portuguese exporters faced considerable uncertainty about whether they would maintain preferential access to Spain. Had they ever lost those preferences Portuguese exports would face an average tariff of 14%.

3.2 Macro Evidence

Following the 1986 accession, Portugal experienced a trade boom. Between 1985 and 1992 *total* Portuguese real exports grew by 90% and imports by about 300%.[24] Figure 1 shows that the share of Portugal's trade with Spain and the EC-10 rose rapidly from 52% in 1985 to 72% by 1992. Remarkably, the initial preferential agreement between the EC-10 and Portugal (agreed in 1972 and implemented by 1977) and Spain and Portugal (early 1980s), which reduced applied tariffs, left their trade share nearly unchanged at about 50% between 1972 and 1985.[25] This is consistent with

[22] The accession did reduce applied protection on agricultural goods. This is why the average preferential tariff faced by Portuguese exporters in the EC was about 2.5% before 1986.

[23] This uncertainty about the elimination of tariffs is clear from a GATT report where one of its members noted that the EFTA-Spain agreement "provided only an expectation that at some point in time the duties and other regulations of commerce would be eliminated but no specific provisions existed in this respect. There was a great difference between an expectation and a specific plan and schedule". "Agreement between the EFTA countries and Spain", Report of the Working Party, L/5405, October 24, 1980, p.3.

[24] Authors' calculations based on data from Pinheiro et al (1997).

[25] We can detect more of an effect during this period if we focus on Portuguese export shares alone, which go from 50% to 62% in this 13 year period. But export growth is faster after the 1986 accession and the EC share in Portugal exports goes up to 73% in only 7 years. The strong increase in trade shares with the EC after 1985 was not merely a

the model if exporters believed that those preferences could be reversed before 1986 but became confident they were secure after accession.

Starting in 1981 we have access to firm-level trade data from the Portuguese census. As a first pass, we check whether the growth in Portuguese exports towards the EC-10 and Spain is also accompanied by a growth in the number of exporting firms and whether either of these can simply be explained by standard trade gravity determinants. More specifically, we test if Portuguese exports to Spain or the EC-10 grew differentially relative to the rest of the world by interacting an EC accession time dummy (=1 for 1986 and subsequent years) with EC-10 and Spain country dummies, and control for standard determinants of trade.[26] The first column of Table 1 shows an increase of 23 log points towards the EC-10 and 115 toward Spain following the accession. In column 2 we use the (ln) number of exporting firms to each market as the dependent variable and find it was significantly higher for EC-10 and Spain after accession.

We also note that the typical new Portuguese exporter to a given market (defined as a firm exporting to a country at t but not $t-1$) is smaller than a continuing firm (about 6 times smaller for Spain and 20 times for the EC-10), which is consistent with our model of heterogenous firms with export entry costs. Moreover, size heterogeneity implies that *if* accession increased the number of entrants but not the average sales of continuing firms, then we should observe a reduction in average sales per firm as we do for the EC-10 in column 3 of Table 1. In the supplementary appendix D.4, we provide further evidence that a large share of the increase in the number of exporters and export volume is due to new entrants (as opposed to fewer exits) and is thus likely to have required considerable entry investments. We also find that even though individual entrants are smaller they had a significant impact on aggregate exports in subsequent years.

3.3 Micro Evidence: Empirical Approach

The macro evidence is suggestive but by itself it does *not* allow us to determine what *caused* the accession effect, whether TPU was reduced and if so how important this was relative to any changes in applied policies. First, factors other than trade policy may have affected exports and firm entry over the long period used in the gravity estimate.[27] Therefore, to minimize the role of other confounding factors we will subsequently focus on the growth between 1985-1987, and

switch away from exporting to other markets. There is strong evidence of trade creation: total real exports in 1993 were almost twice as high as in 1985 (Pinheiro et al., 1997).

[26] We also include bilateral fixed effects to address PTA endogeneity following Baier and Bergstrand (2007). These reduce to importer fixed effects since the only exporter is Portugal. Year effects control for Portuguese productivity and price changes.

[27] Even in the gravity estimate the accession effect is *not* due to two alternative explanations: monetary integration and infrastructure investments. First, the sample we use for the gravity estimate ends in 1990—two years before the signing of the Maastricht Treaty setting out the *timetable* for the Euro and nine years before the exchange rates were irrevocably fixed. Second, even if accession affected exchange rates, these effects are controlled for by including its level and its volatility (see working paper). Portugal received substantial EC support for transport infrastructure, but this *funding* only began in 1989. At best it could have started to reduce trade costs in 1990 and can't explain the large prior trade increases.

note that even over this short period the net growth in the number of exporting firms was on average 25 and 91 log points toward the EC-10 and Spain respectively. Second, to identify a causal effect and estimate the relative importance of applied policies and TPU we must explore the model predictions using firm-level data and theory-consistent policy measures that vary across countries, products and time.

3.3.1 Measuring policy uncertainty

We construct the profit loss measure in (11) by assuming a discrete tariff distribution H. We choose a tractable distribution that covers the main cases present in our data. After a policy shock exporters consider three potential values for the random variable, τ_i: the tariff faced in destination $i = \{\text{EC-10,Spain}\}$. The realized tariff values can (and do) vary over goods but for simplicity we omit those subscripts.

$$\tau_i = \tau_{si}, \quad \Pr(\tau_{si}) = p_{si} \text{ for } s = \{l, m, h\}$$

We take τ_{li} to be 1 for all goods and destinations, i.e. free trade. This captures the industrial goods that Portugal exported to the EC free of ad valorem tariffs both after the accession *and* before it. It also captures the lowest possible tariffs observed at any point. The high tariff, τ_{hi}, is the potential worst case tariff that the Portuguese exporters feared if they lost preferences. In the data we assume this is the MFN rate that the EC (or Spain) applied to GATT members before 1986. Note that τ_{hi} is different for the EC and Spain before 1986 so that will be one source of variation in the uncertainty measure.[28] The intermediate tariff, τ_{mi}, captures *transitional* preferences granted to Portugal before 1986 by Spain in several types of goods and the EC-10 in agricultural goods. The latter preferences were transitional and *not* GATT legal in the absence of a permanent trade agreement. Therefore, although we did observe "medium" tariffs during the mid 80's, the Portuguese exporters likely placed a negligible probability ($p_{mi} \approx 0$) on the event that conditional on a policy shock these would remain since either an agreement would be signed and tariffs would transition to the low state or negotiations would fail and no preferences would remain.

In supplementary appendix D.2 we show that if the tariff was initially high or medium then we can use (11) to derive

$$\omega(\tau_{ti}) - 1 = -p_{hi}\left[1 - (\tau_{ti}/\tau_{hi})^{\sigma}\right] \tag{12}$$

The term in square brackets is the percentage profit reduction conditional on a shock that increases tariffs from τ_{ti} to τ_{hi}, which happens with probability p_{hi}. The same term applies to cases when the initial tariff is low and p_{mi} is negligible.[29] Moreover, the applied tariff may change over time but

[28] This may somewhat underestimate the degree of uncertainty in these goods but seems a reasonable approximation of what the Portuguese exporters may have feared as the worst case scenario before the agreement.

[29] Alternatively, if we consider only a two state world, $s = h, l$, the expression above applies to tariffs with either history. Note that the tariffs are different across markets and industries. In supplementary appendix D.2 we show that if p_m were large then there would be an *additional* term where the high probability and tariff are replaced by the medium ones. We ignore this extra term since, (a) there is no obvious empirical counterpart for the medium

the worst case tariff, τ_{hi}, is constant in the data over the short period of time we analyze. If there is no preferential tariff, as is the case for some goods, then the GATT rate and the applied rate are the same, $\tau_{ti} = \tau_{hi}$ and potential profit reductions are zero. In section 3.4 we further describe how we use information on tariffs and the elasticity of substitution to construct this measure.

We cannot observe the probabilities that exporters place on a worst case scenario, $\gamma_i p_{hi}$. However, we can estimate this as part of the entry equation, provided we assume it is roughly common across industries and countries. The similarity across countries in our application is more reasonable if we restrict our attention to i =Spain, EC-10, the markets to which Portugal secured access. One of our goals is therefore to test if there was a regime change, i.e. if the probability of reversal to the worst case scenario, $r \equiv \gamma p_h$, fell after accession.

3.3.2 Unobserved cutoffs and firm export entry

We do not directly observe if firms have costs below the cutoff but we do observe the number of firms and their export status at the country-product level. Our model focuses on variation in policies over time and across products and the cutoffs we derived are common across some sets of firms. In particular, producers exporting to i face a tariff that does not discriminate by firm but rather by product or industry classification, denoted by V. So, in the model producers of $v \in V$ face the *same* critical cutoff, denoted by c^U_{tiV}. Hence we examine the fraction of exporters in an "industry" V to each country pair. This approach has another advantage: it does not require us to be able to follow specific firms over time, which is important since we are unable to do this between 1985 and 1986 (see Appendix D.3).

The number of firms exporting in V to market i is at least equal to the mass of domestic producers in V, n_{tV}, times the fraction of those firms with costs below the cutoff, $G_V(c^U_{tiV})$. Therefore the observed number of firms, n_{tiV}, can be written as

$$\ln n_{tiV} = \ln G_V(c^U_{tiV}) + \ln n_{tV} + u_{tiV} \tag{13}$$

where u_{tiV} is a random disturbance due to measurement error.[30]

3.3.3 Baseline model

We obtain an estimation equation that identifies key structural parameters as follows. We use a linear Taylor approximation of the cutoff in (9) *around* no uncertainty and use the definitions for

probability term, (b) it would be highly correlated with the high value, and (c) we have good reasons to believe $p_m \approx 0$ given these were transitional tariffs that could not be sustained under GATT rules.

[30]The term can also capture the potential for "legacy" firms: those that survive until period t even though they have costs above c^U_{tiV}. This cannot occur if current conditions are better than in the past, so a sufficient condition to rule out legacy firms is that $c^U_{tiV} \geq \max\{c^U_{TiV} \; \forall \; T < t\}$. In the mid-80's the conditions for Portuguese exporters were improving, as is clear from the observed high entry rates into EC countries. Therefore, we do not think legacy firms pose a significant issue in this particular setting. Furthermore, in Handley and Limão (2012) we argue that our approach and results are robust to certain instances where legacy firms are present.

the uncertainty factor, U, and the profit loss term, $\omega(\tau_{tiV})$, from (12). We substitute the resulting expression for c^U_{tiV} on the RHS of (13). Finally, we assume that productivity follows a Pareto distribution with shape k and minimum value of $1/c_V$, so $G_V\left(c^U_{tiV}\right) = \left(c^U_{tiV}/c_V\right)^k$. We then obtain the following for each t, i and V

$$\ln n_{tiV} = k\left[-r_T \frac{\beta}{1-\beta}\frac{1-(\tau_{tiV}/\tau_{hiV})^{\sigma}}{\sigma-1} - \frac{\sigma}{\sigma-1}\ln\tau_{tiV}\right] \tag{14}$$
$$+\,k\left[\frac{1}{\sigma-1}\ln\frac{A_{iV}}{K_{iV}(1-\beta)} - \ln c_V\right] \quad + \ln n_{tV} + \tilde{u}_{tiV}\ .$$

The term $r_T \equiv \gamma_T p_{Th}$ represents the probability of reversal to the worst case tariff and $\tilde{u}$ captures the random disturbance term and higher order terms from the cutoff's approximation. Recall that at any point in time firms treat the regime as time invariant. In the estimation we want to test if the regime changed after accession, so we include a T subscript in r_T where $T = 0$ denotes parameter values for years pre-accession and $T = 1$ post-accession. We discuss relaxing this and some of the identifying assumptions in the robustness section.[31]

Re-writing (14) in terms of parameters and observables we obtain

$$\ln n_{tiV} = b_{\gamma T}\tilde{\omega}_{tiV} + b_{\tau}\ln\tau_{tiV} + a_{ti} + a_{iV} + a_{tV} + \tilde{u}_{tiV} \qquad \text{for each } t,\ i, V \tag{15}$$

where $\tilde{\omega}_{tiV} \equiv \frac{1-(\tau_{tiV}/\tau_{hiV})^{\sigma}}{\sigma-1}$ captures the uncertainty measure and its impact on entry is estimated by $b_{\gamma T} = -r_T k\beta/(1-\beta)$. The coefficient on the applied tariff is $b_{\tau} = -k\sigma/(\sigma-1)$. To identify the parameters of interest we must be able to control for certain key variables in the model, which we do using various fixed effects. The a_{iV} and a_{tV} terms represent country-industry and industry-time effects that absorb the foreign demand conditions in A_{iV}, the investment cost K_{iV} and other time invariant costs of exporting (e.g. transport or other non-tariff barriers we abstracted from in the theory), the productivity distribution heterogeneity across industries c_V, and industry time varying effects, such as the domestic mass of producers, n_{tV}. The term a_{ti} captures country-year effects that control for any unanticipated log-separable aggregate shocks, e.g. to aggregate expenditure.[32]

We estimate (15) in differences taking a period before and one after the accession.

$$\Delta_t \ln n_{tiV} = b_{\gamma 1}\tilde{\omega}_{1iV} - b_{\gamma 0}\tilde{\omega}_{0iV} + b_{\tau}\Delta_t\ln\tau_{tiV} + a_i + a_V + \Delta_t\tilde{u}_{iV} \quad \text{for each } i, V \tag{16}$$

We focus on $i = \{$EC-10, Spain$\}$ and first address two basic questions:

- Did the agreement generate more entry in industries with higher initial potential loss, $-b_{\gamma 0} > 0$?
- Did the agreement reduce the probability of a worst case scenario: $-b_{\gamma 0} > -b_{\gamma 1}$, which implies $r_1 < r_0$; or even eliminate it: $b_{\gamma 1} = 0$, which implies exporters expect preferences not to be reversed,

[31] The baseline estimation relies on three assumptions to identify the effect of uncertainty: (i) k is common across V (but c_V need not be); (ii) σ is common across V; (iii) r is common across V and i.

[32] Recall from section 2 that A_{iV} is log-separable in importer aggregate expenditure, Y_i. So any unanticipated shocks to Y_i are captured by a_{ti}. In the robustness section we also address the possibility of shocks to the price index.

i.e. $r_1 = 0$?

In order to control for destination and industry shocks our identification relies on differential tariffs and uncertainty that Portuguese exporters in each industry faced in the EC vs. the Spanish market and focus on a short period: 1985-87. Doing so required us to compile a rich data set.

3.4 Data

To estimate (16) we collected and digitized detailed data on trade policy for Spain and the original EC-10 countries before and after the agreement, as described in supplementary appendix D.2. The uncertainty measure we compute varies across industries and members of the agreement. For some industries the policy data are *recorded* at a fine level of disaggregation, so they could potentially be matched to 6-digit NIMEXE classifications for the trade data, which includes over 5000 products (NIMEXE is the predecessor of the Harmonized System). We do not test the model at this disaggregated level for a few reasons. First, the model suggests that we define industries according to a set of characteristics (such as productivity distribution) that is common across a set of firms and broader than the 6-digit level. Second, most of the variation in the policy in this data occurs across industries, rather than within them.[33] Third, an exporter's perception of the worst case tariff scenario is likely to be broader than what is implied by the worst case for a *single* 6-digit good, since he may either export multiple goods and/or fear tariff changes simply because they are reclassified. Thus our measure will reflect the uncertainty of other 6-digit products in the same 2-digit category.[34]

We construct the profit loss uncertainty term as described in section 3.3.1, i.e. taking τ_{hi} for a 6-digit product to be the ad valorem GATT tariff that country i =EC-10 or Spain, had before the agreement.[35] We take τ_{0i} to be the tariff that i actually applied to Portuguese exports in that product before the agreement, namely the pre-existing preferential tariffs they provided to Portugal.[36] To construct the measure in (12) we use an elasticity consistent with the data for these countries ($\sigma = 3$). We then aggregate the ω measure and the applied tariff by taking the simple average over all 6-digit products in each 2-digit industry.

Our firm level data is from the Portuguese census (INE) and we describe it in supplementary appendix D.3. The dependent variable in (16) is the log growth in the number of firms exporting to each country i in each 2-digit industry. Since some firms export more than one product in a

[33] About 80% of the variation in applied tariffs and 75% of variation in the uncertainty measure in exporting to the EC-10 before the agreement are accounted for by differences across 2-digit industries (of which there are 99). Those fractions are lower for Spain but still more than half of the variation is accounted for by cross-industry differences.

[34] If we were to run the model at the 6-digit level there would be a large number of zeroes. Since our estimation equation is in logs we would eventually have to drop those categories, which could be where uncertainty was most important.

[35] If that tariff was not bound in the GATT, then we use the autonomous ad valorem tariff that i applied.

[36] For the EC-10 countries there is no cross-country variation in the policy measures because they have a common trade policy. There is, however, variation in applied tariffs for the EC-10 over time and variation between the EC-10 and Spain in policy within each industry.

2-digit industry to some markets, one of our specifications will use this measure of varieties.

Table 2 provides summary statistics for the tariffs that Portuguese firms faced in Spain in both 1985 and 1987. The average industry in Portugal enjoyed preferential tariffs that were nearly 50% below the tariff Spain levied on the rest of the world. Moreover, this difference is not driven by any one set of goods or industries. Using the measure of profits lost we calculate that if Portugal lost these preferences, the typical exporter would see profits reduced by over 16% per annum. There is also scope for export increases due to a reduction in applied tariffs, which stood at 8% before 1986. Portugal enjoyed lower preferential tariffs from the EC-10 before the agreement but the potential loss in profits measure was high—15% on average. Table 2 also shows that the EC average tariff reduction was small since tariffs in industrial products were zero before accession.

3.5 Baseline Estimates

Table 3 provides estimates of the parameters in (16). Entry is negatively affected by applied tariffs, as predicted by the theory. The coefficient on the uncertainty measure, $-b_{\gamma 0}$, is positive, implying that entry was strongest in the industries that initially faced higher potential profit lost in the worst case scenario.[37]

If protection measures other than tariffs were reduced and more so in industries with higher uncertainty then the results in column 1 would suffer from omitted variable bias. We address this by controlling for changes in "non-tariff barriers" and specific tariffs in columns 3 and 4 respectively. Both have the predicted negative sign but they are insignificant. Neither affects the baseline results for uncertainty and applied ad valorem tariffs. The results are also robust to including other policy measures (in columns 5 and 6), which we discuss below.

The results thus far exclude any uncertainty measure *after* the agreement. This reflects an implicit assumption that the coefficient on that variable is insignificant, i.e. that the agreement was credible and eliminated the preference reversal, $b_{\gamma 1} = 0$. We test this hypothesis in column 2 by including the potential profit loss term evaluated at the post agreement tariffs, $\tilde{\omega}_{1iV}$. We find that this variable has no significant effect, i.e. we can't reject that $r_1 = p_{h1}\gamma_1 = 0$, nor can we reject that the probability of a reversal has fallen, i.e. $r_1 < r_0$. The insignificance of uncertainty *after* the agreement and the fact that the restricted version is preferred by standard information criteria (in the last two rows), leads us to focus on column 1 as the baseline estimates.[38]

What is the relative impact of uncertainty and applied policy changes on entry? One simple measure of this impact is how much variation in entry each of the variables explains. For the full sample we find that a one standard deviation reduction in applied tariffs leads to a 0.14 standard

[37]The policy measures vary across industry and for Spain vs. the EC-10. But they do not vary across the EC-10 so we cluster the standard errors to allow for arbitrary correlation across EC-10 countries within each industry.

[38]The results that we discuss subsequently will be qualitatively unchanged if we included the post uncertainty variable. Moreover, while the magnitude of the tariff and initial uncertainty are somewhat different, their ratio is fairly similar with or without post uncertainty, and as we will see it is that ratio that is key to the quantification.

deviation increase in entry compared with an increase of 0.4 for uncertainty, which is almost 3 times larger. Using the model structure in the next section we will go considerably beyond this in quantifying the impact of each policy and their complementarity. But before doing so we discuss additional potential threats to identification and robustness.

3.6 Identification and Robustness

The baseline estimates address several identification concerns and we summarize the main ones here. We differenced out time invariant country-by-industry effects and explored variation over time and within industry across countries. One potential concern is pre-existing growth trends in specific destination markets or industries (either European demand or Portuguese supply trends in specific industries). To ensure that any such trends do not bias our results we included industry and country effects in the baseline specification in *changes.* These industry and country effects also address other potential sources of bias.[39]

We also examine the robustness of the results along different dimensions, which we summarize here and discuss in detail in the working paper.

1. Alternative policy uncertainty measure. Suppose the relevant measure of uncertainty for exporters is different so our model is misspecified. One alternative policy uncertainty measure is the standard deviation of the tariff faced by Portugal in each industry, i.e. $\Delta\left(stdev \ln \tau_{tiv}\right)$ where $v \in V$. In column 5 of Table 3 we find this variable is insignificant and does not change the value or significance of the theoretically based uncertainty measure.

2. Omitted variable bias. The theory allows the aggregate foreign demand shifter for any given industry to vary across countries, A_{iV}, but not over time. The model can be extended to allow for unanticipated shocks such that $A_{tiV} \equiv Y_{ti}\left(P_{tiV}\right)^{\sigma-1}\mu_V\left(1-\rho\right)\rho^{\sigma-1}$. Such shocks are most plausible in the aggregate expenditure, $Y_{ti,}$ and price index terms. Our differenced estimation equation and country and industries dummies control for shocks to Y_{ti} and some of the potential shocks to P_{tiV}. Thus we are only left with the residual variation in $\Delta \ln P_{tiV}$ at both the country *and* industry level. Portugal is small and therefore not likely to affect the price indices of destination markets. But during this period Spain liberalized against many other countries, which could have affected Spain's price index. If Spain's tariff reductions relative to the rest of the world were correlated with those it made towards Portugal then the baseline estimates would suffer from an omitted variable bias. To test this we collected additional data on changes in Spain's tariffs to the rest of the world to proxy for any residual variation in the price index.[40] In column 6 of Table 3 we find that this variable is

[39] First, they control for possible changes in fixed or sunk costs provided they are industry or destination market specific (e.g. accession could have lowered fixed or sunk costs of entry through streamlining of customs procedures or raised them through additional rules of origin). Second, the accession may have changed the share of intermediates used in Portugal's exports; even if this was differential across industries it would be captured by the industry effects. Third, the accession could have affected exchange rates, but this aggregate shock is addressed by the destination effects.

[40] No such tariff changes were made by the EC-10 against the rest of the world so this variable is zero for the EC-10.

not significant and does not change the baseline effects of either the uncertainty or applied tariff variables. The same is true if we also include all the other applied policy controls in columns 2, 3 and 4.

3. Agriculture. Agricultural products are subject to more non-tariff barriers than others. To some extent these are captured in the baseline by the industry effects. Moreover, including the change in one type of non-tariff barrier does not affect the baseline results (column 3) . If there was an unobserved change in agricultural product barriers *and* it was different across Spain and the EC-10, we would not be able to control for it. But our results are robust to dropping agricultural goods altogether.

4. Elasticity of substitution. The baseline assumes that the *typical* elasticity within any given industry V is 3. This value is consistent with median estimates (across sub industries in each V) calculated using the data in Broda et al. (2008) for Spain and the other EC-10 countries. Moreover, the median elasticity across 2-digit industries (i.e. across V) does not exhibit substantial variation in these countries. There are three industries with somewhat higher elasticities so we verified the baseline results are unchanged if we drop these industries. In addition if we instead use $\sigma = 2$ or 4 we find results similar to the baseline (Table 4).

5. Firm and variety growth. The dependent variable in the baseline is the growth in the number of varieties in an industry to a particular destination. So a firm exporting a single good in an industry V to a market is counted as one variety and if it exports two goods in that industry this will be counted as two. If most firms are single good or each good a firm exports entails a separate sunk cost, then this is reasonable. Otherwise we may want to measure entry by the number of firms exporting in an industry instead of varieties. In Table 2 we see a similar average growth of entry using either the variety or firm definition. Moreover, using the firm definition of entry leads to estimates very similar to those in the baseline, as we show below.

6. Alternative pre-accession period. The baseline results are robust to alternative definitions of the pre-accession period. We verify this by re-estimating the baseline using the growth in varieties between 1987 and the *average* in three years prior to the accession (1983-1985). So the baseline results that use only 1985 as the pre-period are not driven by a "pre-treatment" dip. A similar result holds if we use the firm entry definition. This will be reflected in similar reversal probabilities, as we show below.

In sum, the baseline estimates already address several identification concerns and are robust to other potential concerns.

3.7 Quantification and Counterfactuals

Since the parsimonious baseline specification in Table 3 is closest to the theory and preferred by standard information criteria, we focus on it to estimate the probability of reversal and quantify the importance of uncertainty on entry.

3.7.1 Policy Reversal Estimates

The coefficients for initial uncertainty and tariffs in column 1 of Table 3 map to the following structural parameters: $-b_{\gamma 0} = r_0 k\beta / (1-\beta)$ and $b_\tau = -k\sigma / (\sigma - 1)$. So the estimated probability of reversal before the agreement is

$$\hat{r}_0 = \frac{\widehat{b}_{\gamma 0}}{\widehat{b}_\tau} \frac{\sigma}{\sigma - 1} \frac{1-\beta}{\beta} \tag{17}$$

The first row of Table 4 shows that for the baseline estimate $\hat{r}_0 = 0.39$ when $\beta = 0.85$.[41] Even though we did not constrain this parameter its estimate falls in the theoretically feasible range between zero and one. Moreover, its standard error allows us to reject that it is zero. Recall that $r_0 \equiv p_{h0}\gamma_0$ so the estimate indicates that exporters in 1985 believed that the policy was neither fixed, i.e. $\gamma_0 \neq 0$, nor certain to improve, i.e. $p_{h0} \neq 0$, so the reform was not fully anticipated. The reversal probability is similar if we estimate it using $\sigma = 2$ or 4 as seen in the first column of Table 6.[42]

Given that $p_h\gamma$ captures the *ex-ante* belief of exporters, we can't definitively argue that a particular estimate is too high or low. In order to provide additional quantification it is useful to ask what reform scenarios the estimates are compatible with and whether any seem unreasonable. The baseline reversal estimate of $p_{h0}\gamma_0 = 0.39$ is consistent with two extreme beliefs before the agreement. Either the policy shock is fully anticipated, $\gamma_0 = 1$, and preferences are lost with probability $p_{h0} = 0.39$, or preferences will surely be lost, $p_{h0} = 1$, but the timing of the policy change is uncertain and has probability $\gamma_0 = 0.39$. We can bracket our subsequent quantification using these extremes but will focus on describing the intermediate case that seems more reasonable and where $\hat{\gamma}_0 = 0.62 = \widehat{p_{h0}\gamma_0}^{0.5}$ so the policy shock was likely but not certain, i.e. $\gamma \in (.5,1)$.

3.7.2 Policy Impacts on Export Entry

We now quantify the policy impact on entry and decompose it into three effects:

$$\ln \frac{n(\tau_1, r_1)}{n(\tau_0, r_0)} = \underbrace{\ln \frac{n(\tau_0, r_1)}{n(\tau_0, r_0)}}_{\text{Credibility Effect}} + \underbrace{\ln \frac{n(\tau_1, r_0)}{n(\tau_0, r_0)}}_{\text{Non-credible } \Delta\tau} + \underbrace{\left[\ln \frac{n(\tau_1, r_1)}{n(\tau_1, r_0)} - \ln \frac{n(\tau_0, r_1)}{n(\tau_0, r_0)}\right]}_{\text{Complementarity}} \tag{18}$$

The *credibility* effect measures the growth in entry from eliminating policy uncertainty at the initial tariffs; so it captures the effect from making prior tariff concessions by the EC and Spain

[41] Taking the model literally this implies an annual exogenous death shock probability of 0.15. Other authors have assumed this parameter to be 0.125 (Constantini and Melitz, 2008, p.24) and Portuguese annual firm exit rates from production is about 0.17 (calculated from Quadros de Pessoal), which is an upper bound for the exogenous death shock probability since it includes endogenous exit decisions. If more realistically we model firms discount factor to allow for a positive real interest rate, R, then the discount factor is $\beta = (1-\delta)/(1+R)$ where δ is the annual probability of death and $\beta = .85$ is equivalent to alternative reasonable combinations such as using the average real interest rate for Portugal in 1983-1995, $R = 0.03$, and an annual death shock probability $\delta = 0.125$.

[42] The second column of Table 6 shows that the firm entry definition generates reversal estimates similar to those of the variety entry specifications for each σ.

permanent. The second term measures growth from a counterfactual non-credible agreement that reduced tariffs to τ_1 but left the initial regime unchanged. The complementarity terms measures the effect of tariff reductions made credible in the agreement.

Table 5 shows that non-credible tariff changes have relatively small effects on entry (for computation details see appendix D.1). Overall, the tariff policy impacts led to an increase in entry of 12 log points. Credibility accounts for 65% and complementarity for 16% of the policy impacts. Therefore, if tariffs had been reduced in a non-credible way (i.e. leaving initial uncertainty unchanged) entry would have grown by only about 2.3 log points.[43]

Our structural estimates also allow us to determine what fraction of the credibility effect can be attributed to a pure reduction in risk. To do so we rewrite the credibility effect as follows

$$\underbrace{\ln \frac{n(\tau_0, r_1)}{n(\tau_0, r_0)}}_{\text{Credibility Effect}} = \underbrace{\ln \frac{n(\mathbb{E}(\tau'), r_1)}{n(\mathbb{E}(\tau'), r_0)}}_{\text{Pure Risk Reduction}} + \left[\ln \frac{n(\tau_0, r_1)}{n(\mathbb{E}(\tau'), r_1)} - \ln \frac{n(\tau_0, r_0)}{n(\mathbb{E}(\tau'), r_0)}\right]. \tag{19}$$

The first term is the growth in entry due to credibly setting tariffs permanently at their long-run mean, $\mathbb{E}(\tau')$, so it represents a mean preserving compression in tariffs, as argued in the theory section. If the initial tariff was at the long-run mean, i.e. $\tau_0 = \mathbb{E}(\tau')$, then all of the credibility effect would be accounted for by risk reduction. If the initial tariffs are below the long-run mean, as is the case here, then the agreement will have an additional effect of locking in lower mean tariffs. The latter effect is captured by the term in brackets and is positive because more permanent reductions in tariffs relative to the mean—the first term in brackets—have larger effect on entry than temporary ones—the second term.[44] In Table 8 we see that the credibility effect is about 8 log points and 58% of that is due to a pure risk reduction. For the EC, where most tariffs were already at zero, the pure risk reduction still accounted for more than half of the credibility effect; for Spain, where initial tariffs were closer to the estimated mean, it accounted for 82%.

One final point regarding quantification is the fraction of entry *observed* in the data that the trade policy changes implied by the theory predict. As we see in Table 2, varieties increased by 36 log points in this period so the model accounts for about 1/3 of this. There was a substantial increase in the number of Portuguese producers in this period, 16 log points between 1985-1987 (authors calculations from Quadros de Pessoal). Recall that this is exogenous in our model. Thus we can also ask what is the share of predicted entry probability explained by policy, $\Delta \ln (n_{iV}/n_V)$, we show in Table 5 that trade policy changes explain 61% of this value.[45]

[43] These results are for the overall sample. In the remaining columns we see that a non-credible tariff reduction would have generated only about 40% of the predicted entry for Spain and 11% for the EC.

[44] We estimate the pure risk reduction component using the credibility effect formula (25) in appendix D.1, evaluated at $E(\tau'_{iV}) \approx (1 - \hat{p}_{h0}) \tau_{liV} + \hat{p}_{h0} \tau_{hiV}$. We employ the assumptions about the tariff process described in section 3.3.1 (namely τ_l=1 and $p_m \approx 0$).

[45] The model explains almost all of this increase in the data for the EC and 1/3 in Spain indicating that for the latter other factors were also important, e.g. Spain's high income growth.

3.7.3 Policy Impacts on Export Value

Our focus thus far has been on entry but we can also employ our framework to analyze export values.[46] Industry V exports to country i are given by $R_{tiV} = n_{tV} G\left(c^U_{tiV}\right) \times \bar{R}_{tiV}$ where $\bar{R}_{tiV}$ is the average over firm sales in (1). Using the equilibrium cutoff, c^U_{tiV}, we obtain a gravity equation at the industry level, which we estimate in first differences—between 1987 and 1985—as done for entry.

$$\Delta_t \ln R_{tiV} = B_{\gamma 1}\tilde{\omega}_{1iV} - B_{\gamma 0}\tilde{\omega}_{0iV} + B_\tau \Delta_t \ln \tau_{tiV} + a_i + a_V + \Delta_t \tilde{u}_{iV} \quad \text{for each } i, V, \tag{20}$$

In Appendix B we derive this interpretation for the coefficients: $B_{\gamma t} = -r_t \frac{\beta}{1-\beta}(k - (\sigma - 1))$ and $B_\tau = -\frac{k\sigma}{\sigma-1}$. The estimation confirms the entry results. Both policy variables have the expected sign and are statistically significant, moreover uncertainty has no effect *after* the agreement. The exact parameter estimates for $B_{\gamma t}$ and B_τ themselves are not as interesting as the quantification exercises that they allow us to perform, so we focus on the latter. We derive the implied probability of reversal when $\sigma = 3$ to be $\hat{r}_0 = 0.45$—statistically identical to the estimate obtained using the entry equation. We then use this reversal estimate, assuming $\gamma = \sqrt{0.45}$, and the structure of the model to predict the impact of policy on exports.

Tables 7 and 8 quantify and decompose the policy impacts analogously to what we did for entry (see appendix B for computation details). The total predicted export growth due to policy is given by the sum of the uncertainty removal at initial tariffs, $\ln\left[R_{iV}(r_1)/R_{iV}(r_0)\right]$, and the tariff reduction in the absence of uncertainty. This sum is 34 log points for the EC and Spain combined and accounts for 87% of the export growth in 1985-1987 (net of the increase in the number of portuguese firms, last row of Table 7).[47] The first column of Table 7 shows that credibility accounts for 59% of the predicted 34 log points, which is a large effect. To see this notice that Portugal's exports to the EC-10 and Spain were 14.7% of its GDP in 1987 but would have only been 11.5% if uncertainty had remained at its 1985 level. We also calculate that about half of the credibility effect is attributable to a pure risk reduction (Table 8, column 1).

One final point to note is that if the agreement had taken the form of a non-credible tariff reduction it would have increased exports by less than 10 log points overall. The effect is weaker for the EC-10 (because its tariffs were already low) than for Spain. But even a non-credible

[46] Understanding the determinants of the extensive margin is important for several reasons. First, the extensive margin may dominate the response of trade flows to reductions in trade barriers (Chaney, 2008). Second, conditional on survival, small export market entrants grow over time and make up a significant fraction of export growth, as we show is the case for Portugal in supplementary appendix D.4.

[47] In Table 2 we see the average export growth is 55 log points so if we net out 16 log points (the growth in the mass of Portuguese firms, which is exogenous in the model) we obtain 39 log points. The predicted export growth is higher than the growth due to entry, suggesting that the policy changes also affected average exports. Such an effect would arise naturally if firms could make technology upgrading or capacity building investments after entry. Such upgrading is likely given the large increase in export values for existing firms, that we derived previously. Handley and Limão (2013) extend this model to allow for technology upgrading and show that in this case there is an additional effect of reducing uncertainty on export values.

reduction in Spain's high tariffs would have only lead to about half of the predicted growth to that market. because Portuguese exporters would not have entered as strongly.[48] This contrast in the sources of growth of exports provides another interesting motive to consider both the EC-10 and Spanish case since some recent PTAs may look more like the EC-10 case (e.g. Colombia securing pre-existing preferences received in the U.S. market) and others like Spain's (e.g. Korea obtaining US tariff reductions and securing them).

4 Conclusion

Credibility is an important component of a policy reform; our approach and results show it is possible to measure and quantify the impact of trade policy credibility on firm investment and export decisions. The model delivers clear predictions for how to empirically compute policy uncertainty and estimate its impact. Applying our framework to a particular type of trade agreement we find that, (i) before accession to the EC, Portuguese exporters believed there was a positive probability of losing pre-existing preferences in the EC and Spanish markets but the 1986 accession eliminated this TPU; (ii) the overall trade policy changes account for a considerable share of firm entry and export value in the data and (iii) if the agreement had only reduced the applied tariffs, but not TPU, it would have achieved only 20% of the total predicted growth for entry and less than 30% for total exports. So a substantial fraction of the growth was generated by the elimination of TPU. Moreover, about 65% of the growth in entry is due to securing previous tariff reductions, and more than half of this is attributable to a mean preserving reduction in risk.

These results have several implications. First, it points to one reason why unilateral preferences provided to developing countries by the U.S. and E.U. are not always successful in promoting trade and investment—they are subject to uncertain renewal— and this may change if those preferences are secured through formal PTAs. Second, ex-post analysis of PTAs often finds large trade effects even if applied policies are low. From this it is often inferred that either those applied policies are correlated with other unmeasured but *applied* trade costs that were also reduced, or that their trade elasticity is very high. Our results suggest an alternative explanation: the reduction of TPU. This uncertainty may also help explain the border puzzle: why trade across an international border is considerably smaller than within a country even when trade costs appear similar.

Our framework can be extended and address other important questions. Interesting extensions include the effects of TPU on pricing (under variable markups) and the intensive margin (via technology upgrading). Handley and Limão (2013) incorporate technology upgrading in a 2-country model to identify and quantify the role of the U.S. threat to impose Smoot-Hawley tariffs on China.[49]

[48] The tariff elasticity of exports values under uncertainty is 0.69 of its deterministic counterpart (see Appendix B).

[49] Another interesting extension is to examine the interaction of uncertainty between trade policy and demand conditions, to analyze the role of TPU during the great trade collapse (Carballo, Handley and Limão, 2014).

A Theory Appendix

1. Value Function Solution

To obtain the equilibrium values of exporting, $\Pi_e(c,\tau_t)$, and waiting, $\Pi_w(\tau_t)$ in (7) and (8) we solve the linear system given by (5), (6) and the following two equations contained in them

$$\mathbb{E}\Pi_e(c,\tau') = \mathbb{E}\pi(c,\tau') + \beta\mathbb{E}\Pi_e(c,\tau')$$

$$\mathbb{E}[\Pi_e(c,\tau') \mid \tau' \leq \bar{\tau}] = \mathbb{E}[\pi\left(c,\tau'\right) \mid \tau' \leq \bar{\tau}] + \beta\left[(1-\gamma)\mathbb{E}[\Pi_e(c,\tau') \mid \tau' \leq \bar{\tau}] + \gamma\mathbb{E}\Pi_e(c,\tau')\right].$$

2. Cutoff expression (c_t^U)

We use (7), (8), and (4) to solve for c_t^U as a function of the current tariff

$$c_t^U = \left\{\frac{A}{K}\left[\frac{\tau_t^{-\sigma}}{1-\beta(1-\gamma)} + \frac{\beta\gamma\mathbb{E}(\tau^{-\sigma})}{(1-\beta)[1-\beta(1-\gamma)]} + \frac{\beta\gamma H(\tau_t)[\tau_t^{-\sigma} - \mathbb{E}(\tau^{-\sigma} \mid \tau \leq \tau_t)]}{(1-\beta)[1-\beta(1-\gamma)]}\right]\right\}^{\frac{1}{\sigma-1}} \quad (21)$$

Using this and the definition below for $\omega\left(\tau_t\right)$ we obtain (9) in the text.

3. Profit loss expression and bound ($\omega\left(\tau_t\right) \leq 1$)

$$\begin{aligned}\omega\left(\tau_t\right) &\equiv \left\{\mathbb{E}(\tau^{-\sigma}) + H(\tau_t)[\tau_t^{-\sigma} - \mathbb{E}(\tau^{-\sigma} \mid \tau \leq \tau_t)]\right\}/\tau_t^{-\sigma} \\ &= \left[H(\tau_t)\mathbb{E}(\tau^{-\sigma} \mid \tau \leq \tau_t) + (1-H(\tau_t))\mathbb{E}(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma} - H(\tau_t)\mathbb{E}(\tau^{-\sigma} \mid \tau \leq \tau_t)\right]/\tau_t^{-\sigma} \\ &= \left[(1-H(\tau_t))E(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma}\right]/\tau_t^{-\sigma} \leq 1\end{aligned}$$

where the last inequality follows from $(1-H(\tau_t))\mathbb{E}(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma} \leq \tau_t^{-\sigma}$ and the fact that the LHS is a weighted average of two terms, one equal to $\tau_t^{-\sigma}$ and the other equal to $\mathbb{E}(\tau^{-\sigma} \mid \tau > \tau_t)$, which is less than $\tau_t^{-\sigma}$. When the current tariff is at the maximum of the support of $H(\tau)$ such that $\tau_t = \tau_h$, then the difference in brackets and the term $(1-H(\tau_t))$ are both zero.

4. Entry monotonically decreasing in γ

To show this we simply take the derivative of (21) to obtain

$$-\left.\frac{d\ln c_t^U}{d\gamma}\right|_{\tau_t} = -\frac{\beta}{1-\beta(1-\gamma)}\frac{1-\beta}{1-\beta\left(1-\gamma\omega\left(\tau_t\right)\right)}\frac{\omega\left(\tau_t\right)-1}{\sigma-1} \geq 0 \quad (22)$$

which is non-negative given $\omega\left(\tau_t\right) \leq 1$.

5. Equivalence of γ increase and mean preserving spread if $\tau_t = \mathbb{E}\left(\tau'\right)$

In section 2.5 we note that if $\tau_t = \mathbb{E}\left(\tau'\right)$ then increases in γ induce a mean preserving *spread* of the tariff, so they correspond to a pure increase in risk. To see this, denote the tariff distribution conditional on the current tariff and policy regime by $F\left(\tau_{t+1}|\tau_t,\gamma\right)$ and note that if $\tau_t = \mathbb{E}\left(\tau'\right)$ then $\mathbb{E}_t\left(\tau_{t+n}|\tau_t\right) = \mathbb{E}\left(\tau'\right)$ for any γ since $\mathbb{E}_t\left(\tau_{t+n}|\tau_t\right) = (1-\gamma)^n\tau_t + (1-(1-\gamma)^n)\mathbb{E}\left(\tau'\right)$. So change sin γ leave any n period ahead mean unchanged. To show the mean preserving spread result of increasing γ we derive $F\left(\tau_{t+1}|\tau_t,\gamma\right)$ as a function of the current tariff τ_t and the unconditional distribution $H(\tau)$, after a shock γ, as follows

$$F\left(\tau_{t+1}|\tau_t,\gamma\right) = \Pr(\tau_{t+1} < \tau) = \begin{cases} \gamma H(\tau) \text{ if } \tau < \tau_t \\ 1-\gamma+\gamma H(\tau) \text{ if } \tau \geq \tau_t \end{cases}$$

When $\tau_t = \mathbb{E}\left(\tau'\right)$ ($\equiv \int_{\tau_l}^{\tau_h} \tau dH(\tau)$) then an increase in γ to γ^{hi} induces a new distribution $F\left(\tau_{t+1}|\tau_t,\gamma^{hi}\right)$ with the same mean, as shown above. We now show that $\int_{\tau_l}^{\tau} F\left(\tau_{t+1}|\tau_t,\gamma^{hi}\right) d\tau \geq \int_{\tau_l}^{\tau} F\left(\tau_{t+1}|\tau_t,\gamma\right) d\tau$ for any $\tau \leq \tau_h$ so $F\left(\tau_{t+1}|\tau_t,\gamma\right)$ second order stochastically dominates $F\left(\tau_{t+1}|\tau_t,\gamma^{hi}\right)$. This clearly holds when $\tau < \tau_t$ since $\int_{\tau_l}^{\tau} \gamma^{hi} H(\tau)d\tau \geq \int_{\tau_l}^{\tau} \gamma H(\tau)d\tau$. To prove the condition, for $\tau \geq \tau_t$ we note that $\int_{\tau_l}^{\tau} F\left(\tau_{t+1}|\tau_t,\gamma\right) d\tau = (1-\gamma)[\tau-\tau_l] + \int_{\tau_l}^{\tau} \gamma H(\tau)d\tau$ is increasing in γ when $\tau_\tau = \mathbb{E}\left(\tau'\right)$

$$\begin{aligned}
\frac{\partial}{\partial\gamma}\left[\int_{\tau_l}^{\tau} F\left(\tau_{t+1}|\tau_t,\gamma\right) d\tau\right] &= -[\tau - \mathbb{E}\left(\tau'\right)] + \int_{\tau_l}^{\tau} H(x)d\tau \\
&= -\tau + \mathbb{E}\left(\tau'\right) + [H(\tau)\tau - H(\tau_l)\tau_l] - \int_{\tau_l}^{\tau} \tau dH(x) \\
&= -\tau(1-H(\tau)) + \mathbb{E}\left(\tau'\right) - H(\tau_l)\tau_l - (H(\tau)-H(\tau_l))\mathbb{E}(\tau_{t+1} \mid \tau_l < \tau_{t+1} < \tau)) \\
&= (1-H(\tau))[\mathbb{E}(\tau_{t+1} \mid \tau_{t+1} > \tau) - \tau] \geq 0.
\end{aligned}$$

The second line uses integration by parts. The third simplifies and uses the definition of conditional mean. The last line follows by writing the unconditional mean as a weighted average of the conditional means $\mathbb{E}\left(\tau'\right) = H(\tau_l)\tau_l + (H(\tau)-H(\tau_l))\mathbb{E}(\tau_{t+1} \mid \tau_l < \tau_{t+1} < \tau)) + (1-H(\tau))(\mathbb{E}(\tau_{t+1} \mid \tau_{t+1} > \tau)$, where we allow for a mixed distribution H to be continuous or discrete at the end points, and canceling terms.

6. Current tariff impact on profit loss ($d\omega(\tau_t)/d\tau_t \geq 0$)

As τ_t increases, the profit lost from being hit with a shock to a higher tariff is reduced so $\frac{d\omega_t}{d\tau_t} > 0$

$$\begin{aligned}
\frac{d\omega(\tau_t)}{d\tau_t} &= [-\tau_t^{-\sigma}h(\tau_t) + h(\tau_t)\tau_t^{-\sigma} - \sigma H(\tau_t)\tau_t^{-\sigma-1}]/\tau_t^{-\sigma} + [(1-H(\tau_t))\mathbb{E}(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma}](\sigma\tau^{\sigma-1}) \\
&= [-\sigma H(\tau_t)\tau_t^{-1}] + \sigma\tau^{\sigma-1}[(1-H(\tau_t))\mathbb{E}(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma}] \\
&= \sigma\tau^{\sigma-1}[-H(\tau_t)\tau_t^{-\sigma} + (1-H(\tau_t))\mathbb{E}(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma}] \\
&= \sigma[(1-H(\tau_t))\mathbb{E}(\tau^{-\sigma} \mid \tau \geq \tau_t)]/\tau^{1-\sigma} \geq 0
\end{aligned}$$

In semi-elasticity terms, this becomes

$$\frac{d\omega(\tau_t)}{d\ln\tau_t} = \frac{\sigma[(1-H(\tau_t))\mathbb{E}(\tau^{-\sigma} \mid \tau \geq \tau_t)]}{\tau^{-\sigma}} \in [0,\sigma] \tag{23}$$

This implies that as the current tariff τ_t increases, the proportional gap between the deterministic and uncertain cutoff narrows. We can see that that if $\tau_t = \tau_h$ the derivative goes to zero. Then $\frac{d\ln c_t^U}{d\ln\tau_t} = -\frac{\sigma}{\sigma-1}$ and the elasticity of the cutoff under uncertainty evaluated at the tariff maximum equals the elasticity at the deterministic cutoff.

7. Attenuated tariff impact on entry under uncertainty: $\frac{d\ln c_t^U}{d\ln\tau_t} \Big/ \frac{d\ln c_t^D}{d\ln\tau_t} \in [0,1]$

Using the expression for c_t^U from the text, we log differentiate and derive

$$\begin{aligned}
\frac{d\ln c_t^U}{d\ln\tau_t} &= \frac{d\ln c_t^D}{d\ln\tau_t} + \frac{d\ln U_t}{d\ln\tau_t} \\
\frac{d\ln c_t^U}{d\ln\tau_t} &= -\frac{\sigma}{\sigma-1} + \frac{1}{\sigma-1}\frac{\beta\gamma}{1-\beta(1-\gamma\omega)}\frac{d\omega_t}{d\ln\tau_t} \\
\frac{d\ln c_t^U}{d\ln\tau_t} / \frac{d\ln c_t^D}{d\ln\tau_t} &= 1 - \frac{\beta\gamma}{1-\beta(1-\gamma\omega)}\frac{d\omega_t}{d\ln\tau_t}\frac{1}{\sigma}
\end{aligned} \tag{24}$$

As we show in (23) $\frac{d\omega_t}{d\ln\tau_t}\frac{1}{\sigma} \in [0,1]$, so entry is less responsive to tariff changes under uncertainty except at the two limiting cases when $\gamma = 0$ (deterministic τ) and when τ_t is at the maximum.

B Estimation and Quantification Appendix

In this appendix we describe how to implement the decomposition of the policy effects in Tables 5, 7 and 8 and derive the estimating equation for total exports presented in the text.

1. Decomposing and quantifying the policy impacts on entry

To estimate the average credibility effect component of equation (18), we use the relationship between n and the cutoff in (13) and the equilibrium cutoff in (9) to obtain

$$\begin{aligned}
\ln\frac{n(\tau_{0iV}, r_1)}{n(\tau_{0iV}, r_0)} &= k\left[\ln U(\tau_{0iV}, r_1) - \ln U(\tau_{0iV}, r_0)\right] \\
&= -\frac{k}{\sigma-1}\ln\left[\frac{1-\beta(1-\gamma_0\omega(\tau_{0iV}))}{1-\beta(1-\gamma_0)}\right]
\end{aligned} \tag{25}$$

where the second line uses $U(\tau_{0iV}, r_1 = 0) = 1$, as suggested by the baseline estimates. The expression in brackets can be rewritten in terms of the estimated parameters, $-b_{\gamma 0}$ and b_τ, the data $\tilde{\omega}_{0iV}$ and a given γ such as $\hat{\gamma}_0 = 0.62$, which we can then average over the observations. Doing so we obtain the credibility effect of about 8 log points in Table 4. Given there were about 40428 country×(firm-product) pairs in total in the sample this predicts an additional 3167 pairs by 1987.

We also compute the counterfactual effect of a non-credible agreement that reduced only applied tariffs. First, because τ also affects the uncertainty term, U, we must determine how attenuated the entry elasticity is under uncertainty relative to no uncertainty, i.e. $\frac{d\ln n(\tau_0,r_0)}{d\ln\tau} \Big/ \frac{d\ln n(\tau_0,r=0)}{d\ln\tau}$. From (13) we see that this is equal to the ratio of the cutoff elasticities, $\frac{d\ln c^U}{d\ln\tau} / \frac{d\ln c^D}{d\ln\tau}$, which we derive in the appendix, eq. (24) derived above. Using the definition of ω in (12) in (24) we obtain $\frac{d\omega_0}{d\ln\tau}\frac{1}{\sigma} = p_{h0}\left(\frac{\tau_{0iV}}{\tau_{hVi}}\right)^\sigma$, which we combine with estimated parameters and data. In Table 4 we show that this factor is 0.56, so under the initial uncertainty the same tariff reductions would have generated only about half as much entry than if uncertainty was absent (so only slightly over 2 log points overall, 1 for EC and 11 for Spain).[50]

[50] The result is not sensitive to the choice of σ. The magnitude of the attenuation does depend on how much of the reversal, $p\gamma$, is due to the arrival shock. The effect is bounded by the extremes: if the arrival shock is very likely ($\gamma = 1$) the attenuation is 0.71, if it is unlikely ($\gamma = 0.39$) then the attenuation is 0.35.

The impact of applied policy on entry in the absence of uncertainty, the term $\ln \frac{n(\tau_1,r_1)}{n(\tau_0,r_1)}$ in (18), is given by multiplying the deterministic entry elasticity, b_τ, and the change in applied tariffs. This yields an overall entry growth of about 4 log points (2 for the EC-10 and 20 for Spain since the latter had larger tariff reductions). Therefore the total predicted entry due to removal of applied tariffs and uncertainty is about 12 log points (4+8), slightly lower for EC-10 (10) and higher for Spain (28), as we report in Table 5.

2. Aggregation to industry exports

The total export value to a given country in an industry V is

$$R_{tiV} = \left[n_{tV} G_V \left(c^U_{tiV}\right)\right] \times \bar{R}_{tiV} \tag{26}$$

where average sales $\bar{R}_{tiV} \equiv A_{iV}\sigma\tau_{tiV}^{-\sigma}\tilde{c}$ are obtained by averaging (1) over all exporting firms and when $G_V\left(c^U_{tiV}\right) = \left(c^U_{tiV}/c_V\right)^k$ we have $\tilde{c} = \frac{k}{k-(\sigma-1)}\left(c^U_{tiV}\right)^{1-\sigma}$. Using the cutoff expression and simplifying we obtain

$$\begin{aligned} \ln R_{tiV} &= (k-\sigma+1)\ln c^U_{tiV} - k\ln c_V + \ln n_{tV} + \ln\frac{k}{k-(\sigma-1)} - \sigma\ln\tau_{tiV} + \ln A_{iV} + \ln\sigma \qquad (27) \\ &= \frac{k-\sigma+1}{\sigma-1}\ln\left[\frac{1-\beta\left(1-\gamma\omega\left(\tau_{tiv}\right)\right)}{1-\beta\left(1-\gamma\right)}\right] - \frac{k\sigma}{\sigma-1}\ln\tau_{tiV} + \alpha_{tiV} \end{aligned}$$

where $\alpha_{tiV} = \frac{k-\sigma+1}{\sigma-1}\ln\frac{A_{iV}}{K_{iV}(1-\beta)} + \ln n_{tV} - k\ln c_V + \ln\frac{k}{k-(\sigma-1)} + \ln A_{iV} + \ln\sigma$. We take a first order approximation of the uncertainty term around $\gamma = 0$ and substitute that and the constructed measure $\tilde{\omega}_{tiV}$ in (27) to obtain

$$\ln R_{tiV} = B_{\gamma T}\tilde{\omega}_{tiV} + B_\tau \ln\tau_{tiV} + a_{ti} + a_{iV} + a_{tV} + \tilde{u}_{tiV} \quad \text{for each } t,\ i, V \tag{28}$$

where $B_{\gamma T} = -r_T\frac{\beta}{1-\beta}(k-(\sigma-1))$, $B_\tau = -\frac{k\sigma}{\sigma-1}$ and the a_x terms capture all the terms in α_{tiV} defined above. If we then difference this equation we obtain (20).

3. Decomposing and quantifying the policy impacts on industry exports

The impact of removing uncertainty at initial tariffs on exports used in Table 7 is

$$\ln R_{tiV}|_{\gamma_1} - \ln R_{tiV}|_{\gamma_0} = -\frac{k-\sigma+1}{\sigma-1}\ln\left[\frac{1-\beta\left(1-\gamma_0\omega\left(\tau_{0iV}\right)\right)}{1-\beta\left(1-\gamma_0\right)}\right]. \tag{29}$$

The total impact of applied tariff changes under uncertainty is

$$\frac{\partial\ln R}{\partial\ln\tau} = -\frac{k\sigma}{\sigma-1}\left[1 - \frac{k-\sigma+1}{k}\frac{\beta\gamma}{1-\beta\left(1-\gamma\omega\right)}\frac{\partial\omega_t}{\partial\ln\tau_t}\frac{1}{\sigma}\right]. \tag{30}$$

The leading term is the full elasticity of total exports to tariff changes and the term in brackets is the attenuation factor, which is equal to $\frac{\partial\ln R}{\partial\ln\tau}|_{\gamma_0>0}/\frac{\partial\ln R}{\partial\ln\tau}|_{\gamma_1=0}$, and is reported in Table 7. As above for the entry equation, we can compute the attenuation factor from estimated parameters of the export equation and data. We compute the effect of the tariff reduction at initial uncertainty in Table 7 by multiplying the attenuation factor by the change in tariffs, as if uncertainty remained high.

C Data Appendix

Variable	Description	Source
Aggregate Regressions (Table 1, 1981-1990):		
Exports (ln)	Nominal value of exports in euro of all goods to country i in year t.	a
Number of Firms Exporting (ln)	# of uniquely identified shippers with positive exports to i in year t.	a
Exports per Firm (ln)	ln(Exports$_{it}$/Number of firms$_{it}$)	a
Real Importer GDP (ln)	Country i, year t in billions of importer currency.	b
Importer Price Index (ln)	ln(nominal GDP)-ln(real GDP) in local currency	b
Annual Exchange Rate	Simple average of ln monthly rate, where latter is defined as ln((escudo/importer currency)/200.482). The fixed conversion factor from esc to euro is 200.482 and plays no role in the regressions.	b
Ex-Rate Volatility(ln)	Standard deviation of log monthly changes in the year.	b
Firm and policy data in estimates of Tables 2-8:		
Change in Number of Firms(ln)	ln(# firms exporting to i in V, 1987)-ln(# firms exporting to i in V, 1985) where i is an EC-11 country and V corresponds to a NIMEXE 2-digit industry	a
Change in Number of Firm-Varieties(ln)	ln(# varieties exported to i in V, 1987) - ln(# varieties exported to i in V, 1985), "varieties"defined as distinct 8-digit NIMEXE products exported by each firm	a
Change in exports(ln)	ln(export value to i in V, 1987) - ln(export value to i in V, 1985)	a
Pre Tariff (GATT)	$\ln\tau$ where τ is 1+advalorem rate at product level that GATT members faced in Spain or EC-10, which is then averaged to Nimexe 2-digit industry.	c,d
Pre Tariff (Portugal)	$\ln\tau$ where τ is 1+advalorem rate at product level Portugal faced in Spain or EC-10, which is then averaged to Nimexe 2-digit industry.	c,d
Post Tariff (Portugal)	$\ln\tau$ for immediate post agreement period that Portugal faced in Spain or EC-10, constructed as described in previous section.	c,d
Applied Tariff Standard Deviation Change	Δstd($\ln\tau$) where the standard deviation is over tariffs Portugal faced in each Nimexe-2 industry; the change is between the pre and post tariff.	f
Uncertainty	Proportional reduction in per period profits if the tariff faced by an exporter reverts from the preferential tariff received prior to accession (Pre Tariff above) to the tariff received by all non-preferential partners (i.e. the GATT member tariff).	g
NTM Share Change	Change in coverage ratio measured by fraction of products in 2-digit industry subject to a specific tariff or other NTB.	c,d
Specific Tariff Share Change	Difference in the share of lines in 2-digit industry with specific tariffs between post and pre-agreement period.	f
Price Index Proxy Change	Difference in Spain's external tariff, ln(1+ad valorem rate), between post and pre-agreement period	f
Other data (Figures and text):		
Import & Export to GDP ratios	Referenced in text	h
Trade Shares	Fig. 1, IMF Direction of Trade Statistics	
Export price index (ln)	1985 base chain price index for exports.	h
Employment	Number of employees	i
Firm identifier (NPC)	Unique code to match firms between 1981-1985. Portuguese customs changed this code in 1986 and it is consistent for 1986 onwards but not between 1985 and 1986	INE
New exporter in year t	Firm exporting to a market at t but not in t-1	i
Gross entry rate in year t	(Total # new exporters in t)/(# exporters t-1).	i
Gross exit rate in year t	(# exporters with positive exports in t-1 and none in t)/(# exporters with positive exports in t)	i

Notes (additional details in supplementary appendix):
a. Authors' calculations based on INE data
b. IMF International Financial Statistics (IFS), volatility uses monthly data.
c. Spain: International Customs Journal, Spain, No. 24, 16th Edition, 1984.
d. EC: Official Journal, L342, 31.12.1979, p. 1-382.
e. Articles of Accession, Official Journal L 302 , 15/11/1985.
f. Authors' calculations based on tariff schedules
g. Authors' calculations using equation (12) with $\sigma = 3$ in the baseline regressions.
h. Authors' calculations from in Pinheiro et al (1997). Indices are yearly price deflators of export goods to all destinations.
i. Authors' calculations using trade data matched to firm employment data (Quadros Pessoal) by INE.

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Figures and Tables

Figure 1

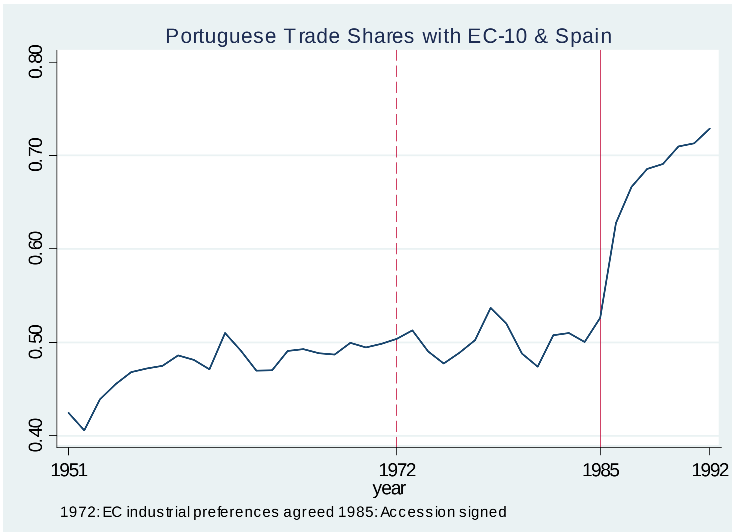

Table 1: Portuguese Export Growth Margins 1981-1990

	(1)	(2)	(3)
Dependent variable (ln):	Exports	Number of Firms	Exports/firm
EC10×Post_86	0.232***	0.451***	-0.219***
	[0.0829]	[0.0411]	[0.0710]
Spain×Post_86	1.146***	1.159***	-0.0129
	[0.199]	[0.132]	[0.113]
Real Imp. GDP (ln)	1.045***	0.598***	0.447*
	[0.306]	[0.137]	[0.258]
Imp. Price Index (ln)	0.167**	0.0185	0.148**
	[0.0776]	[0.0374]	[0.0655]
Exchange rate (ln)	0.211***	-0.00118	0.212***
	[0.0763]	[0.0365]	[0.0653]
Observations	1305	1305	1305
Adj R2	0.92	0.97	0.75

Notes:
Includes dummies for country and year. Robust standard errors in brackets. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$ Sample: Aggregate values to each country of destination. Column 3 is the logarithm of total exports over number of firms. For variable definitions and sources see Appendix C. See Supplementary Appendix, Table D4 for summary stats.

Table 2: Summary statistics for firm-level baseline regressions

	Total	Spain	EC-10
Change in Number of Firms*	33.0 (55.1)	91.1 (62.6)	24.7 (48.7)
Change in Number of Varieties*	35.7 (60.7)	101 (69.1)	26.4 (53.3)
Change in Exports*	55.3 (157)	135 (150)	43.9 (155)
Pre Tariff** (Portugal)	3.13 (5.66)	7.89 (5.10)	2.45 (5.40)
Pre Tariff** (GATT)	8.67 (5.14)	14.1 (7.75)	7.95 (4.20)
Post Tariff** (Portugal)	1.74 (3.91)	1.33 (3.51)	1.79 (3.96)
Tariff Change** (Portugal)	-1.39 (2.90)	-6.56 (4.78)	-0.66 (1.44)
Applied Tariff Stand. Dev. Change***	-0.64 (1.25)	-2.86 (1.86)	0.33 (0.69)
Price Index Proxy Change***	-0.19 (0.88)	-1.52 (2.06)	0.00 (0.00)
NTM Share Change***	-2.32 (10.9)	-18.66 (25.4)	0.00 (0.00)
Specific Tariff Share Change***	-0.37 (2.80)	-3.01 (7.45)	0.00 (0.00)
Proportion of Profits Lost if Preference Reversed			
Initial Uncertainty	15.5 (10.9)	16.0 (9.52)	15.4 (11.1)
Post Uncertainty	18.8 (10.8)	29.4 (15.0)	17.3 (9.08)
Observations	731	91	640

Notes:
Sample means and standard deviations (in parentheses), all multiplied by 100.
* $100 \times \Delta \ln(x)$ where $x = \{\text{firms, varieties, exports}\}$.
**$100 \times \ln(1+t)$ where t is the advalorem rate; "Pre tariff" is evaluated in 1985 (pre-accession); one measures Portugal's preferential rate and the other tariffs faced by GATT members; "Post Tariff" is the 1987 (post-accession) tariff faced by Portugal; "Tariff Change" is a simple difference.
*** See Appendix C for sources and additional details. Profit loss:$1-(\tau_{0V}/\tau_{hV})^{\sigma}$ (assuming $\sigma=3$). We normalize the loss measures in regressions by dividing it by $\sigma-1$.

Table 3: Firm-variety entry growth into EC-10 and Spain (by industry)

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable :	Change in (ln) Number of Firm-varieties					
Initial Uncertainty ($-b_{\gamma 0} > 0$)	4.399** [1.772]	5.626** [2.756]	4.301** [1.810]	4.431** [1.788]	4.351** [1.839]	4.752** [1.854]
Applied Tariff Change ($b_\tau < 0$)	-3.006** [1.260]	-4.273* [2.271]	-3.072** [1.266]	-2.919** [1.247]	-3.113** [1.291]	-3.520*** [1.260]
Post Uncertainty ($b_{\gamma 1} \leq 0$)		-1.51 [3.277]				
NTM Share Change			-0.166 [0.256]			
Specific Tariff Share Change				-0.579 [1.034]		
Applied Tariff SD Change					0.468 [4.144]	
Price Index Proxy Change						1.946 [2.066]
Observations	731	731	731	731	731	731
R-squared	0.471	0.471	0.472	0.472	0.471	0.471
No. of parameters	101	102	102	102	102	102
AIC	1083	1085	1084	1084	1085	1084
BIC	1551	1558	1557	1557	1558	1558

Notes:
Structural parameters and expected sign in parentheses below regressor names. All specifications include country and industry effects. Clustered standard errors in brackets (industry×EC-10). *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Sample: Spain and EC 10 countries, 1987-1985. Assumes $\sigma = 3$. AIC and BIC denote Akaike and Bayes Information Criterion. See Table 2 for summary statistics.

Table 4: Reversal, attenuation and entry estimates

Value of $\sigma =$	2	3	4
Probability of Reversal (standard error)	0.36 (0.16)	0.39 (0.17)	0.42 (0.18)
Tariff elasticity attenuation factor	0.56	0.56	0.55
Entry growth, uncertainty removal (mean ω)	0.08	0.08	0.08
Entry growth, uncertainty removal (min ω)	0.24	0.24	0.24

Notes:
We use the initial uncertainty estimate from Table 3 in calculations. Probability of reversal= $\frac{\hat{b}_{\gamma 0}}{\hat{b}_\tau} \frac{\sigma}{\sigma-1} \frac{1-\beta}{\beta}$, see Section 3.7.1, with s.e. obtained using delta method ($\beta = 0.85$). Conditional on p_h, the attenuation and theoretical uncertainty term U can be derived using regression estimates independently of β assumptions. We take $p_h = (\text{Pr. reversal})^{0.5}$ and compute attenuation and entry growth at mean ω. The attenuation factor is the ratio of the entry elasticity to tariff changes at initial uncertainty relative to no uncertainty, see equation (24). We compute entry due to uncertainty removal as the log difference in the number of entrants at post vs. initial uncertainty using $k \times [\ln(U_1) - \ln(U_0)]$ and assume post-uncertainty is removed, $U_1 = 1$.

Table 5: Entry counterfactuals and quantification

	Total	Spain	EC
Predicted entry effect of tariff and uncertainty reductions	0.12	0.28	0.10
Share of predicted effect due to:			
Uncertainty removal (at initial tariffs)	0.65	0.28	0.80
Tariff reduction (at initial uncertainty)	0.19	0.40	0.11
Complementarity	0.16	0.32	0.09
Share of predicted entry probability explained by policy	0.61	0.33	1.03

Notes:
We use the initial uncertainty estimate from Table 3 in calculations (but results are similar if $\sigma = 2,\ 4$). Predicted average entry probability is the sum tariffs reductions, $b_\tau \Delta \ln \tau$, and removal of uncertainty at initial tariffs. See Appendix B for further details. Entry growth from uncertainty removal is the log difference in the number of entrants at post vs. initial uncertainty using $k \times [\ln(U_1) - \ln(U_0)]$ (we assume post-uncertainty is removed, $U_1 = 1$). Counterfactual shares of predicted entry hold initial tariffs and uncertainty fixed, respectively. The complementarity share captures the remaining entry growth from simultaneously reducing tariffs and uncertainty. The share of predicted entry probability explained by policy is the ratio of policy predictions relative to the one observed in the data.

Table 6: Probability of reversal robustness

Sample	1987-1985		1987-pre mean	
Dependent variable (ln), no. of:	varieties	firms	varieties	firms
Value of $\sigma = 2$	0.36	0.36	0.27	0.26
3	0.39	0.39	0.29	0.28
4	0.42	0.41	0.31	0.30

Notes:
Assumes $\beta = 0.85$, for any other value of β, simply divide by 0.15/0.85 and multiply by the new $(1 - \beta)/\beta$. The 1987-1985 sample uses the growth between 1987 and 1985, which is the baseline. The 1987-pre mean sample uses the growth between 1987 and the average of the three years before the agreement: 83, 84, 85.

Table 7: Total exports counterfactuals and quantification

	Total	Spain	EC
Predicted export effect of tariff and uncertainty reductions	0.34	0.85	0.26
Share of predicted total exports due to:			
Uncertainty removal (at initial tariffs)	0.59	0.23	0.75
Tariff reduction (at initial uncertainty)	0.28	0.53	0.17
Complementarity	0.13	0.24	0.08
Share of predicted total exports explained by policy	0.87	0.72	0.96

Notes:
Calculations use an initial uncertainty estimate from the total export regression where $\sigma = 3$ and the probability of reversal is estimated at 0.45, see section 3.7.3 and Appendix B for details (results are similar if $\sigma = 2,\ 4$). Average predicted total exports are the sum of the tariff reduction, $B_\tau \Delta \ln \tau$, and the effect of uncertainty removal at initial tariffs. Total export growth from uncertainty removal is the log difference in exports at post vs. initial uncertainty using $(k - \sigma + 1) \times [\ln(U_1) - \ln(U_0)]$ (we assume post-uncertainty is removed, $U_1 = 1$). Counterfactual shares of predicted exports hold initial tariffs and uncertainty fixed, respectively. The complementarity share captures the remaining export value growth from simultaneously reducing tariffs and uncertainty. The share of total export growth explained by policy is the ratio of policy predictions relative to exports observed in the data net of the aggregate growth in number of Portuguese firms.

Table 8: Credibility effect from securing pre-existing reforms

	Total	Spain	EC
Entry growth:			
Total effect (ln points)	7.8	8.7	7.7
Share from risk reduction (at long-run mean)	0.58	0.82	0.54
Export growth:			
Total effect (ln points)	19.7	22.0	19.6
Share from risk reduction (at long-run mean)	0.51	0.72	0.47

Notes:
Calculations use an initial uncertainty estimate from the entry or total export regressions where $\sigma = 3$ and the probability of reversal is estimated at $r_0 = 0.39$ and $r_0 = 0.45$, respectively. Total growth from uncertainty removal is the log difference in entry or exports at post vs. initial uncertainty using k or $(k - \sigma + 1)$ times $[\ln(U_1) - \ln(U_0)]$ evaluated at the pre-period applied (τ_0) and GATT tariff (τ_h) means from Table 2 (we assume post-uncertainty is removed, $U_1 = 1$). The counterfactual shares from risk reduction evaluate $\ln(U_0)$ while holding applied tariffs at $\tau_0 = E(\tau)$ to obtain a mean preserving compression of the tariff distribution. We compute $E(\tau)$ from our estimates of $\gamma_0 p_h$. See section 3.7.2 for details of the calculation (results are similar if $\sigma = 2,\ 4$).

D Supplementary Appendix (FOR ONLINE PUBLICATION ONLY)

D.1 Theory

1.. First-order Cutoff Approximation

We take a first-order Taylor approximation of $\ln c^U$ around $\tau_t = \tau_0$ and $\gamma = 0$.

$$\begin{aligned}\ln c^U(\gamma, \tau_t) &= \ln\left(c_t^D \times U_t\right) \\ &\approx \ln c^D(\ln \tau_0, \gamma = 0) + + (\ln \tau_t - \ln \tau_0)\frac{\partial \ln c_t^D \mid_{(\tau_0, \gamma=0)}}{\partial \ln \tau} + (\ln \tau_t - \ln \tau_0)\frac{\partial \ln U_t \mid_{(\tau_0, \gamma=0)}}{\partial \ln \tau} \\ &+ \ln U(\ln \tau_0, \gamma = 0) + \gamma\frac{\partial \ln c_t^D \mid_{(\tau_0, \gamma=0)}}{\partial \gamma} + \gamma\frac{\partial \ln U_t \mid_{(\tau_0, \gamma=0)}}{\partial \gamma}\end{aligned}$$

Using the definition for c_t^D, equation (22), noting that $\partial \ln U_t / \partial \ln \tau|_{\gamma=0} = 0$ and $\partial \ln U_t / \partial \gamma \mid_{\tau_0, \gamma=0} = \frac{\beta}{1-\beta}\frac{\omega-1}{\sigma-1}$, and simplifying we obtain the expression the first order approximation used for the regression specification in the main text

$$\ln c_t^U \mid_{\tau_t=\tau_0, \gamma=0} \approx \gamma\frac{\beta}{1-\beta}\frac{\omega(\tau_0)-1}{\sigma-1} - \frac{\sigma}{\sigma-1}\ln \tau_t + \frac{1}{\sigma-1}\ln\frac{A}{K(1-\beta)} \tag{D1}$$

2. Impact of a simple mean preserving spread in $H(\tau)$ on entry cutoff

Claim: For any current tariff $\tau_t \in (\tau_l, \tau_h)$, the entry cutoff is higher under $H(\tau)$ than under a simple mean preserving spread of it, $F(\tau)$, with the same support and $F(\tau_t) = H(\tau_t)$.

Proof: From the cutoff condition we see that it will be higher under H if, conditional on a bad shock, expected profits are higher under H than F, i.e. if $\frac{1}{1-H(\tau_t)}\int_{\tau_t}^{\tau_h}\pi(c,\tau)dH(\tau) > \frac{1}{1-F(\tau_t)}\int_{\tau_t}^{\tau_h}\pi(c,\tau)dF(\tau)$. By construction $F(\tau_t) = H(\tau_t)$ so we need to show $\int_{\tau_t}^{\tau_h}\pi(c,\tau)dH(\tau) - \int_{\tau_t}^{\tau_h}\pi(c,\tau)dF(\tau) > 0$. Using integration by parts we have $\int_{\tau_t}^{\tau_h}\pi(c,\tau)dX(\tau) = [\pi(c,\tau_h)X(\tau_h) - \pi(c,\tau_t)X(\tau_t)] - \int_{\tau_t}^{\tau_h}\pi'(\tau)X(\tau)d\tau$ for $X = H, F$. The term in brackets is identical for $X = H, F$ since by construction $H(\tau_h) = F(\tau_h) = 1$ and $F(\tau_t) = H(\tau_t)$. Thus we obtain

$$\int_{\tau_t}^{\tau_h}\pi(c,\tau)dH(\tau) - \int_{\tau_t}^{\tau_h}\pi(c,\tau)dF(\tau) = \int_{\tau_t}^{\tau_h}\pi'(\tau)\left[F(\tau) - H(\tau)\right]d\tau > 0.$$

The inequality follows because (i) if F is a simple mean preserving spread of H and $F(\tau_t) = H(\tau_t)$ then $F(\tau) < H(\tau)$ for $\tau \in (\tau_t, \tau_h)$ and (ii) $\pi'(\tau) < 0$.

D.2 Policy Data

1. Pre-accession policy data

The earliest trade data for Portugal is from 1981 and the closest full EC trade policy schedule before then is for 1980 (Official Journal L 342, 31.12.1979, p. 1–382). This, and the fact that EC applied tariffs to Portugal in industrial goods were the ones set in the 1977 agreement, and thus remained in place until 1985, lead us to initially digitize and use the 1980 schedule.[51] The 1980 schedule already reflects some of the EC multilateral tariff bindings negotiated in the Tokyo Round.

[51] While our baseline results only use data for 1985 and 1987 to isolate the effect of the agreement in 1986, we also planned and ran robustness tests that include earlier years.

However, some of these bindings, which we use to construct our uncertainty measure, continued to be reduced over a period of time.[52] Therefore, if the worst case scenario for Portuguese exporters between 1981-1985 was the EC binding then it may have entailed a lower tariff than that implied by the 1980 binding. Even for those goods where the binding was falling the 1980 binding may still be the appropriate one to capture the exporter expectations we model if for example the exporters did not immediately update their beliefs about the tariff distribution.

We obtained the 1984 trade policy schedule for Spain. This schedule was published by the International Customs Tariff Bureau in the *International Customs Journal*. We believe this was the only full schedule published in the 1980s for Spain and it contains Spain's preferences relative to Portugal and the EEC as well as its policy relative to the rest of the world. The documentation we found implies that Spain's preferential tariffs for Portugal remained unchanged between 1984 and 1985 because the EFTA-Spain agreement that regulated these had reached a phase requiring additional negotiations of indeterminate length.

2. Post-accession policy data

To construct the tariff profile faced by Portugal immediately after the agreement we applied the concessions schedule in the Articles of Accession, Protocol 3 for Spain (Official Journal L 302 , 15/11/1985 P. 0410) and Article 243 for the EC (Official Journal L 302 , 15/11/1985 P. 0094). These imply staged reductions of 12.5% per year for Spain and 14.2% for EC-10 with some exceptions for certain goods and industries. Portugal and Spain also harmonized their tariff with respect to the rest of the world to match the EC common external tariff.

3. Applied Protection and Uncertainty Measures

The schedules for the EC and Spain were manually keyed into digital format at the tariff line level by a firm specialized in data entry. We performed a number of checks to ensure that the quality of the entry and kept track of the few tariff lines with various combinations of minimum and maximum tariffs, specific tariffs and seasonal tariffs.[53] We then applied preference margins for the EFTA-Spain and EC-Portugal agreements to compute the applied tariff faced by Portuguese exporters in 1985. We applied the staged reductions of the Articles of Accession to these schedules for the EC and Spain to compute the 1987 tariff profile. These digitized schedules yield our tariff line measures of applied tariffs in 1985 and 1987. There are about 9500 tariff lines for Spain and 6500 lines for the EC in any particular year. Finally, we digitized a set of pre- and post-accession NTMs applied by Spain at the 4-digit industry level based on accession documentation submitted to the GATT.[54]

4. Concordance and Aggregation

We constructed our tariff panels using the Brussels Tariff Nomenclature to maintain consistency between published schedules and the preference margins stipulated in pre- and post-accession agreements. However, our firm level data are classified by Nimexe so we map each BTN code to a 6-digit Nimexe code using a time-consistent 6-digit Nimexe. We constructed the concordance by digitizing the EC's official concordance between the BTN and Nimexe.[55] To further maintain

[52]"Implementation of MTN concessions: Note by the secretariat, revision" TAR/W/8/Rev.3, October 15, 1981

[53]For example, Spain levies an ad valorem tariff of 14% on product 66.01-A-I "Umbrellas and sunshades: Covered with fabrics of silk or man-made fibres" subject to a minimum specific tariff of 75 pesetas each. We use the ad valorem tariff as our tariff line applied measure and track the presence of the minimum tariff in an indicator variable.

[54]See "List of Non-Tariff Restrictive Measures Applied by Portugal and Spain before and after their Accession" L/5936/Add.5, 5 March 1987.

[55]See "Commission Regulation(EEC) No 3840/86 amending the nomenclature of goods for the external trade statiscts of the Community and statistcs of trade between Member States (NIMEXE)" (Official Journal L368,

time-consistency, our concordance allows for changes the Nimexe system over the sample period. We tracked these yearly changes according to schedules found in the Eurostat publication External Trade Nomenclature of Goods, Volume 5 (1990). When there are multiple BTN codes mapped to single Nimexe code, we average within the Nimexe code. The same schedules give us the pre- and post- accession worst case tariff used to compute the uncertainty measure as described in the main text. We aggregate by industry up to the 2-digit Nimexe level by taking the arithmetic mean of tariffs and our uncertainty measures. Within each industry, we keep track of detailed tariff line information by computing the shares of tariff lines with complex and specific tariffs and use these as additional controls in the robustness checks.[56]

5. Implementation of Tariff Uncertainty in Discrete Case

To construct the empirical measure of $\omega(\tau_t)$ we consider a discrete probability distribution for tariffs. We then ask, given that a policy shock above the current trigger τ_t arrives, what is expected value of the proportional loss in profits? This quantity is summarized by $\omega(\tau_t) - 1$, which we now compute for a two- and three-state tariff process relevant to our empirical implementation.

Two State Tariff Distribution: High, Low

Initial State ($\tau_T = \tau_s$)	Probability (p_s)	$\omega(\tau_T = \tau_s) - 1$
τ_h	p_h	0
τ_l	$1 - p_h$	$-p_h\left(\tau_l^{-\sigma} - \tau_h^{-\sigma}\right)/\tau_l^{-\sigma}$

In the two state case, any firm with an entry trigger $\tau_t \geq \tau_h$would enter when the tariff is in the high state. The likelihood of a shock to trade policy leading to a worse outcome is zero. As was the case with a general continuous distribution, the cutoffs in the deterministic and uncertain model will coincide at the maximum. In the low state, $\omega(\tau_t) - 1$ is nonzero and less than unity. In the estimation, we construct the observable counterpart to the $\omega(\tau_t) - 1$ from tariff data and assumptions on σ.

Three State Tariff Distribution: High, Medium and Low

Initial State ($\tau_T = \tau_s$)	Probability (p_s)	$\omega(\tau_T = \tau_s) - 1$
τ_h	p_h	0
τ_m	p_m	$-p_h\left(\tau_m^{-\sigma} - \tau_h^{-\sigma}\right)/\tau_m^{-\sigma}$
τ_l	$1 - p_m - p_h$	$-\left[(p_m + p_h)\left(\tau_l^{-\sigma} - E(\tau^{-\sigma} \mid \tau > \tau_l)\right)\right]/\tau_l^{-\sigma}$
		$= -\sum_{s=m,h}[p_s(\tau_l^{-\sigma} - \tau_s^{-\sigma})]/\tau_l^{-\sigma}$

The three state distribution is slightly more involved, but makes it clear how to generalize to many discrete states. We argue in the empirical section that Portugal had "medium" preferential tariffs with respect to Spain by 1983 of an indefinite nature due to the EFTA-Spain agreement. If $p_m \longrightarrow 0$, then we see that the measures in the second and third row coincide with our empirical implementation for the EC and Spain.

29/12/1986).

[56] Our estimation method requires an industry level net entry dependent variable, but this is not the only reason to aggregate tariff line policy data over the sample period. From 1980 to 1987 the Brussels Tariff Nomenclature (BTN) and Nimexe classifications are updated annually, but our tariff schedules reflect the classification in place at the time of implementation. Spain, Portugal and the EC further differentiate tariffs within BTN categories. BTN and Nimexe are time-consistent and equivalent at the 2-digit level.

D.3 Firm and Aggregate Data

Our firm level data is from the Portuguese census (INE). We use the transaction level trade data available for the period 1981-1990 from customs declaration forms processed by INE. Since the 1981-1987 trade data had not previously been used we did several basic exercises to check their accuracy. We found no law establishing minimum value thresholds for filling out the customs forms in this period. There are no discontinuities at low values in the shipment value distribution. We confirmed that the aggregate yearly values of both imports an exports matched those reported by the official INE printed publication "Estatisticas do Comercio Externo" for several years. INE converts data for all years into euros at a rate of 200.482 esc/euro even before the euro was implemented.

1. Firm identifiers

Evidence on new vs. continuing firms makes use of the shipper's identifier variable (labelled NPC) to determine if it is a new or existing exporter to a market. INE reports that this is a unique firm identifier after 1986 and it is in fact used to match trade data to employment and other firm-level data collected by INE in recent years in other work. There is also a unique identifier in 1981-1985 but so far neither INE nor Portuguese customs have been able to provide a correspondence that would allow linking specific firms between 1985 and 1986. Given that the pre-1986 data had not previously been used we requested INE to confirm with Portuguese customs that pre-1986 identifiers were unique and allowed us to track firms in that period, which they did. We further investigated this by calculating statistics by NPC in each year (e.g. industry of modal product exported, # products, # destinations, total shipment value and weight, etc) and verifying they were highly correlated in adjacent years, e.g. the elasticity of total export values by NPC between 1985 and 1984 is one, similarly for other variables. Moreover, these relationships were identical to those found when comparing adjacent years in the post-1986 data where the NPC identifier was known to be unique.

2. Destination country

To ensure that country codes are consistent over time we used the official list of changes in trade partners published yearly in the "Estatisticas do Comercio Externo". When a country splits, the code for the "larger" unit (e.g. Russia) is the same as the existing (e.g. USSR) and a new code is created for others (e.g. Ukraine). When a country merges (e.g. Germany) we assign the same code as the largest of the existing (West Germany) and drop the other (East Germany).

3. Summary Statistics

In addition to the firm-level summary statistics for the baseline regressions in Table 2, we report summary statistics in Table C4 for the macro evidence gravity regressions in Table 1.

D.4 Additional evidence for new vs. continuing exporters

In this appendix we (1) discuss theoretical and empirical motives for focusing on net or gross entry; (2) provide additional evidence on the contribution of new firms for export growth post 1986 and (3) provide analysis complementing the aggregate gravity evidence in section 3 with information for new vs. continuing exporters.

D.4.1 Model implications for net entry

The model in section 2 has implications for both net and gross entry. We could potentially consider them separately, but the nature of the data and entry/exit processes may not always allow us to distinguish between them separately. The central prediction of interest is that certain reductions in uncertainty lower the cost threshold and, all else equal, imply larger numbers of firms exporting. This larger number can be due to two effects

1. Entry of "new" firms: Firms that previously did not export and would not have entered the market in the absence of this uncertainty reduction but now do so.
2. Re-entry or non-exit of "existing" firms: if a firm is hit by a shock that leads it to exit then if it is still below the threshold it will re-enter but if it is above the threshold it will not. When the threshold falls with uncertainty this firm is now more likely to re-enter or not exit in the first place.

To test the central prediction (lower uncertainty leading to higher number of exporters) we must construct an appropriate counterfactual. In Table 1 we considered the number of firms exporting to a preferential market relative to those exporting to the rest of the world before vs. after the PTA while controlling for aggregate determinants of entry. To decompose the effect of uncertainty into new vs. existing firms using the exact same approach and time period we would require data on a consistent set of firm identifiers over the full period. These identifiers are not available in our data for 1985-1986. We do have consistent identifiers for a few months prior to when the accession is implemented (early 1986) and after accession, which we explore below. But even if we had consistent firm identifiers over the full period we cannot observe the "death" shock and may therefore not observe the exit at all (e.g. if a firm suffers a cost increase that would make it exit if the cutoff were unchanged but not if the cutoff cost increased). Therefore we emphasize net entry in our results, because if we focused *only* on gross entry (e.g. those by firms never before in that market) we could miss an important effect of uncertainty.

We provide some additional motivation for the model and rule out some alternative explanations for the increase in firm export entry in the aggregate data. One concern in particular is whether the agreement increased re-entry (or lowered exit) and thus could have lead to a higher number of firms even if the threshold had not changed.

D.4.2 New vs. continuing firms: definitions and stylized facts

To fix ideas we decompose the number of firms exporting to a particular market into continuers (C_t), defined as those that exported at t and $t-1$, and entrants (E_t), those that export at t but not $t-1$. The total number of exporters, n_t, is therefore

$$n_t = C_t + E_t$$

Since $C_t = (1 - exit_t) \times n_{t-1}$ (where $exit_t$ is the exit rate between t and $t-1$, i.e. $\frac{\text{number of exits}}{n_{t-1}}$), we can write the net entry rate as

$$\frac{n_t}{n_{t-1}} - 1 = entry_t - exit_t$$

When we examine the raw data we find that average yearly exit rates in the period after agreement (87-90) are similar to pre-agreement for Spain (about 0.35) and actually increase for the EC-10. Growth in the number of firms in these markets appears to be driven by gross entry. Some direct evidence for this is provided in Figure D1 comparing the yearly number of entering firms into different markets. There is a differentially larger effect for Spain and EC-10. The EC and ROW levels and trend prior to the agreement are very similar but after the agreement the EC had on average almost 800 additional new exporters per year. Spain started from a lower level than the ROW but ended up near the same level and the differential increase in new exporters is more than 750. The effect is more pronounced when compared to a single large market such as the US.

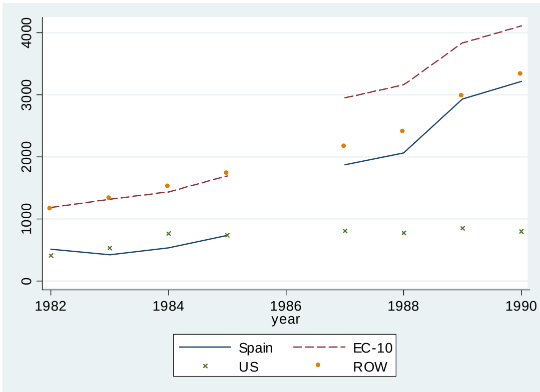

Figure D1: Number of New Firms Exporting to Preferential and Other Markets

Figure D1 also indicates that the increase in the number of exporting firms can't be explained solely by continuing firms choosing to stay in EC markets following accession. Our approach is to identify a counterfactual with the cohort of firms we observed exporting to the market just before the agreement was implemented, or shortly thereafter (since some of the potential new entrants may have waited until implementation to decide to make the fixed cost investment). Because the consistent firm identifier starts in January 1986 we use the cohort of firms that were exporting in 1986. If the agreement did not lead to immediate entry then the change in the stock of firms relative to the 1986 cohort provides a reasonable measure of the stock of "new" entrants. But if some entry already occurred, as the large increase in net entry in 1986 suggests, then this will be an underestimate.

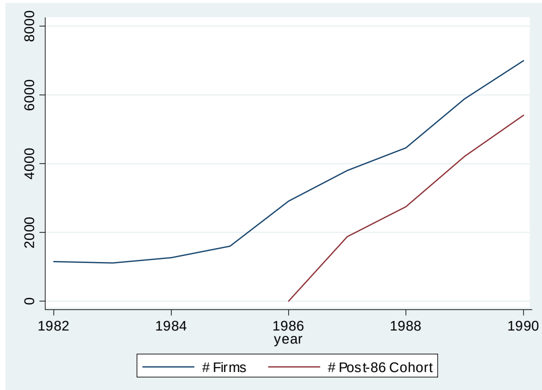

Figure D2: Number of Total vs
Post-Accession Exporters to Spain

Figure D2 shows the total number of firms and those in the post-86 cohort. The latter are zero in 1986 by definition and increase quickly, clearly driving the growth in the total number of firms. The narrowing difference between the lines indicates a decline in continuing firms from the 86 cohort, which implies that re-entry for that initial cohort is insufficient to offset exit. This suggests that the agreement effect is not simply on re-entry or non-exit of firms present before the agreement but on entry of new firms.

D.4.3 Contribution of new firms to exports: 1986-1990

Table D1 uses the raw data to further quantify the aggregate importance of new Portuguese exporters to EC-10 and Spain after accession. Since we are interested in capturing the more immediate effects of the agreement and initially do not control for other factors we focus on a period close to the accession date: 1986-1990. Between these years alone, real exports to the world grew by 50% driven in large part by exports to Spain and the EC, which account for 23% and 59% of that growth. Because these calculations start in 1986, when we can track specific firms, they miss the considerable growth that already took place relative to 1985 (about 7%) and was fully driven by the EC-10 and Spain (real exports to the ROW actually fell in 1986). We find that 46% of the increase in exports between 1986 and 1990 to Spain and 62% to EC-10 is from "new" firms, i.e. those that did not export to those markets in 1986.

We view these numbers as a lower bound of the contribution of firms that entered after the agreement since they exclude the many firms that entered in 1986. Despite this data limitation, we can provide some additional criteria to identify firms that are "likely" new exporters in 1986. The agreement only began implementation in March 1986. So some firms may have waited until that date or later in the same year before starting to export: either because the actual tariff reductions did not start until March (in case of Spain and some EC agricultural products) or because they wanted to be certain that the agreement would in fact be implemented. In fact, we observe that the typical first month of shipment across all firms to Spain in 1986 is May, which is two months later than the median in previous years. For EC-10 countries the median increased about 1 month in 1986. This suggests that we can use a firm's first shipment to these markets in 1986 as a way to help identify when it enters. Using this information we augment the "new exporters" category to include those that export in 1987 or later and also those firms that exported in the second half of 86 but not the first. This criterion identifies about 532 firms in the case of Spain but it is important to note that they only account for about 2% of exports to Spain in 1986. For the EC the criterion typically identifies about 200 firms. Using this criterion the share of new firms in export growth between 1986 and 1990 goes up to 54% for Spain and 73% for the EC, as shown in Table D1.

When making year to year comparisons one potential issue is that if one of the groups fares particularly badly in the final year then the decomposition may be misleading. We account for this by averaging exports for each cohort over the years 1987-1990 and calculating the difference relative to 1986. Doing so generates results similar to Table D1. In sum, we find new exporting firms accounted for a significant fraction of export growth to Spain and EC between 1986 and 1990.

D.4.4 Decomposing the effect of accession: new vs. continuing exporters

The raw data in Table D1 shows that entrants account for a substantial share of export growth between 1986 and 1990. However, this is consistent with different motives for new entry: (a) it is common to all export markets; (b) entry relative to Spain and EC is already present before the agreement, or; (c) it is generated by other determinants, e.g. income and price changes. We rule out these possibilities in Table D2, in which we report the gravity regression coefficients due to accession for exports of continuing vs. new entrants, total number of new entrants, and average annual exports from new entrants. There are two basic changes relative to the sample in Table 1 in the text. First, the new sample excludes countries with zero exports of existing firms in 1981 or 1986 or zero exports of new firms, so the control group is in some ways more similar to the "treatment". Second, we exclude the initial years, 1981 and 1986, since we use them to define firm status.[57]

As we can see in column 3 of Table D2 there is a positive significant effect of the agreement on total exports of new firms for both Spain and EC-10. Decomposing this effect into the number of new firms (column 4) and their average sales (5), we see that it is the former that increases exports. In fact, we see that the average sales of new firms post agreement are lower, which provides evidence that the agreement changed the entry threshold making it easier for smaller firms to export.

To decompose the impact of the agreement we use the estimates to predict the average change in export value to a market implied by the agreement for new firms (column 3) and existing ones (column 2) for each year 87-90, deflate it and average them. The sum of these two predictions is in Table D3: over 500 million euro for Spain and over 600 for EC-10 of additional exports on average per year after 1986. New firms account for about 30% of this predicted change to either of them.[58]

The last two columns of Table D3 show a similar decomposition for the number of firms using the estimates in Table D2 (column 4) and an analogous specification for existing firms (not shown). The predicted average increase in the number of firms due to the agreement is about 3500 for Spain and over 7200 for all the EC-10 countries combined. A large part of this change is from new firms (0.68 for Spain and 0.78 for the EC), as we report in section 3.

[57]We must necessarily do when we use the post-initial year definition. We do the same for comparability purposes when employing alternative definitions of new entrants that may include some exports in the initial year. Neither of these changes to the sample has much effect on the aggregate export specification (not shown), which looks similar to the one for the full sample in column 1.

[58]More specifically, denoting y_{itf} as the log nominal value of exports to i at t for type of firm f we obtain the predicted change as $\exp \mathbf{E}\left(y_{itf} | PTA_i = 1\right) - \exp \mathbf{E}\left(y_{itf} | PTA_i = 0\right)$ for each t>1986 and i=Spain or EC, f=new or existing. We then average over the years and deflate using a 1985 based export price index.

Supplementary Appendix Tables (FOR ONLINE PUBLICATION ONLY)

Table D1: Change in real exports from 1986 to 1990 and shares of "new" firms (million euro, 1985 base)

	Total	Share of post-86 firms[a]	Share of post mid-86 firms[b]
Spain	785	0.46	0.54
EC	1966	0.62	0.73

Notes:

(a) Fraction of change in exports accounted for by firms that did not already export in 1986.

(b) Fraction of change in exports accounted for by firms that did not already export in 1986 *and* those that exported only in the second half of 86 and also first half of 1987.

We treat export to an EC-10 country as a new entry even if the firm exports to a different member.

Table D2: Growth Margins of existing vs. new firms

	(1)	(2)	(3)	(4)	(5)
Sample:	All Firms	Existing firms		New firms[a]	
Dependent variable (ln)	*Exports*	*Exports*	*Exports*	*Number*	*Exports /firm*
EC10*Post_86	0.232***	0.145	0.302**	0.506***	-0.204*
	[0.0829]	[0.0945]	[0.141]	[0.0539]	[0.115]
Spain*Post_86	1.146***	1.173***	1.178***	1.338***	-0.159
	[0.199]	[0.172]	[0.288]	[0.131]	[0.214]
Real Imp. GDP (ln)	1.045***	0.463	1.792***	0.757***	1.036**
	[0.306]	[0.332]	[0.487]	[0.183]	[0.410]
Imp. Price Index (ln)	0.167**	0.235**	0.19	0.0884	0.102
	[0.0776]	[0.107]	[0.124]	[0.0573]	[0.0967]
Exchange rate (ln)	0.211***	0.311***	0.204*	0.0522	0.151*
	[0.0763]	[0.112]	[0.116]	[0.0535]	[0.0914]
Observations	1,305	900	900	900	900
Adj R2	0.92	0.943	0.87	0.964	0.605

Notes: Heteroskedasticity robust s.e. in brackets. *** p<0.01, ** p<0.05, * p<0.1. All specifications include year and country effects. Sample (a) in columns (3)-(5) defined as follows. For 1987-1990: firms that did not export in 1986 and those that exported only in the second half of 86 and also first half of 1987. For 1982-1985: firms that did not export in 1981 and those that exported only in the second half of 81 and also first half of 1982. Existing firms are all others exporting to that market. The larger sample size in the first columns is due to including 1981 and 1986 as well as export destinations served only by existing or new firms.

Table D3: Predicted yearly change in exports and number of firms due to accession & shares of "new" firms

	Exports (million euro)		*Number of firms*	
	Average change	Share "new" firms[a]	Average change	Share new firms[a]
Spain	544	0.27	3546	0.78
EC	654	0.31	7207	0.68

(a) Fraction of predicted change in value of export or number of firms of those that did not already export in 1986 *and* those that exported only in the second half of 86 and also first half of 1987.

(b) Export values in million euro (1985 base)

(c) We treat export to an EC-10 country as a new entry even if the firm exports to a different member. The predicted change refers to the sum over all such countries.

Table D4: Summary Statistics for Gravity Regressions

Variable	Mean	SD	Obs
Exports (ln)	13.780	3.065	1305
Number of Firms Exporting (ln)	3.478	1.986	1305
Exports per Firm (ln)	10.303	1.488	1305
Real Importer GDP (ln, bil. importer currenc)	5.161	3.366	1305
Importer Price Index (ln)	-1.412	1.839	1305
Exchange Rate (ln)	-2.087	3.209	1305