EXPORTING UNDER TRADE POLICY UNCERTAINTY: THEORY AND EVIDENCE

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Abstract

I provide novel evidence for the impact of trade policy uncertainty on exporters. In a dynamic, heterogeneous firms model, trade policy uncertainty will delay the entry of exporters into new markets and make them less responsive to applied tariff reductions. Policy instruments that reduce or eliminate uncertainty, such as binding trade policy commitments at the WTO, increase entry. The predictions are tested on disaggregated, product-level Australian imports with model-consistent measures of uncertainty. The estimates show that growth of exporter-product varieties would have been 7% lower between 1993 and 2001 without the binding commitments implemented after the WTO was formed in 1996. If Australia reduced all its tariffs and bindings to zero, more than half of predicted product growth is accounted by by removing uncertainty. These results illuminate and quantify an important new channel for trade creation.

JEL Codes: D8, F1, F4.

Keywords: policy uncertainty, trade, World Trade Organization, bindings

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1 Introduction

Policy commitment and credibility are often thought to be extremely important for inducing economic agents to make investments, particularly when they entail large irreversible costs. Trade policy is one area where commitment and credibility are potentially very important (Limão and Maggi, 2013; Maggi and Rodriguez-Clare, 1998; Tang and Wei, 2009). A founding principle of the World Trade Organization (WTO) is to establish predictability of trade policy[1]. Despite this objective, a substantial share of the trade between WTO members takes place under flexible trade policy regimes where trade barriers are subject to change. Whether this creates policy uncertainty that has quantifiable impacts on trade has not been well understood. This is partly because most research focuses on trade policy in static, deterministic frameworks. But more importantly, evidence measuring the importance of policy uncertainty in the trade context is limited. I provide novel evidence that when trade policy is uncertain, multilateral policy commitments are an important channel of gains from trade agreements.

Even though the potential for large scale "trade wars" currently seems remote, trade policy uncertainty is pervasive in the world trade system. For example, in the wake of the financial crisis in 2008, leaders of the G-20 repeatedly pledged not to "...repeat the historic mistakes of protectionism of previous eras."[2] Such assurances were necessary because there exists wide scope for protectionism even *within* the WTO. Members make enforceable commitments not to raise applied tariffs above maximum binding constraints.[3] These "bindings" are presently well above applied tariffs in some countries. Over 30 percent of the tariff lines of WTO members could be increased unilaterally without providing compensation to affected trade partners (Bchir et al., 2005). Brazil, for example, could raise tariffs from an average of 11.5 to 36.2 percent; Indonesia from 6.7 to 35.6 percent, and; the average developing country from 8 to 28 percent (Messerlin, 2008). In short, the worst case scenario if governments were to backslide into protectionism, yet not violate any WTO rules is large.[4]

Securing multilateral commitments to eschew 1930s era protectionism was a founding principle of the General Agreement on Tariffs and Trade (GATT), the precursor of the WTO. The 1948 GATT charter explicitly states "binding against increase of low duties or of duty-free treatment shall *in principle* be recognized as a concession equivalent in value to the substantial reduction of high duties or the elimination of

[1]Under the principle "Predictability: through binding and transparency" the WTO explains that "Sometimes, promising not to raise a trade barrier can be as important as lowering one, because the promise gives businesses a clearer view of their future opportunities" http://www.wto.org/english/thewto_e/whatis_e/tif_e/fact2_e.htm (accessed January 20, 2014)

[2]http://www.londonsummit.gov.uk/en/summit-aims/summit-communique/ (accessed November 9, 2010).

[3]A country that violated its bindings would have to provide compensation to affected trade partners or face WTO sanctioned retaliatory tariffs.

[4]There are also other ways to increase protection within the WTO that can and have been used in the past such as anti-dumping cases, invoking special safeguard tariffs, or raising other non-tariff barriers.

tariff preferences."[5] The use of tariff bindings and the existence of gaps between applied and binding rates are a feature of optimal trade agreements (Amador and Bagwell, 2011).[6] But in practice, the principle that constraints on future policy could be as valuable as applied tariff concessions has never been widely accepted or quantified; the trade off continues to be a considerable source of controversy in multilateral negotiations (Evenett, 2007; Mattoo and Subramanian, 2008).

My main contribution is to empirically examine the impact of tariff binding commitments, which lie at the heart of the GATT/WTO, on trade and export market entry. Little is known about effect of bindings on trade because most empirical research has focused on aggregate flows or applied protection. The cross-country study in Rose (2004), for example, questions whether there are any tangible benefits to WTO membership. In contrast, Subramanian and Wei (2007) find the WTO does promote trade when controlling for differential rates of liberalization and access to other preferences. Firm-level evidence in Buono and Lalanne (2012) finds weak extensive margin effects for the tariff changes induced by the Uruguay Round when the WTO was created. However, they do not control for the change in binding commitments induced by the round, instead focusing only on applied concessions.

The channel explored here is uncertainty over trade policy. Dixit's (1989) seminal paper on firm entry and exit under uncertainty shows that when sunk market entry costs are combined with uncertainty over future conditions there may be an option value of waiting to invest. New exporters face both of these elements: evidence suggests there are large sunk cost of entry (cf. Roberts and Tybout, 1997) and there is substantial uncertainty over trade policy. Existing models of policy uncertainty have been largely theoretical. Francois and Martin (2004) provide simulation evidence that by truncating the distribution of tariffs, WTO bindings on agricultural products reduced tariff volatility and raised welfare. In an independent theory piece, Sala et al. (2010) model the impact of bindings in a real options framework but don't provide empirical evidence; they solve the model numerically and then assess the impact of changes in tariffs and bindings for different parameterizations. I provide a bridge from the theory to evidence by extending the tractable, heterogeneous firms trade model in Handley and Limão (2012) to encompass binding tariff commitments. Prospective entrants compare the value of beginning to export today versus waiting. On the margin, the present value of the difference between exporting and waiting reflects only the potential for "bad news" and this leads firms to delay entry. Bindings reduce uncertainty by constraining the range of observable tariffs and limiting losses in the worst case scenario.

I use the model to empirically quantify the policy uncertainty that arises through gaps between applied tariffs and bindings for Australia. I write the model in terms of a latent variable capturing the value of entry

[5] Emphasis added. United Nations Conference on Trade and Employment, Final Act and Related Documents, Interim Commission for the International Trade Organization, April 1948, p. 31

[6] Beshkar et al. (2011) extend the theory and find empirical support for its predictions with WTO binding commitments.

and estimate a linear probability model of observing trade in a disaggregated product as a measure of firm entry. This approach is complementary to Handley and Limão (2012), who focus on how preferential trade agreements can reduce trade policy uncertainty and induce entry rather than how variation in bindings at the WTO effect entry. Using firm-level data, they show that the reduction in uncertainty following Portugal's accession to the European Community explains a substantial share of net entry into EC markets. The method used here is novel for two reasons: first, I am able to use the observable levels of tariff bindings to *test* for the impact of uncertainty with *product-level* data when the standard deterministic model is nested as the null hypothesis; second, the uncertainty measures can be directly controlled by policy so I can use the estimated model to quantify the relative impact of reducing applied protection versus the impact of reducing binding commitments.

The empirical method requires detailed product level trade data and corresponding data on applied and bound tariffs for a single importer. I focus on the role of Australian trade policy for exports to Australia from 1991-2001. High quality and detailed data on products and tariffs are available during this period and, more importantly, there is wide variation across products in binding commitments in both the cross-section and through time. Bound rates range from zero to as high at as 55 percent. On average, Australia's MFN tariff is 4.5 log points during this period. Binding commitments are twice as high, at 9.4 log points. This provides an ideal setting to identify whether or not binding tariff commitments matter for entry and to quantify their role in multilateral liberalization.[7] As described in Section 2, other aspects of Australian trade policy raise issues of uncertainty that are hardly unique to this application.

I find that lowering bindings, while holding applied tariffs fixed, brings the entry decision forward by reducing the incentive to delay investment. The estimates indicate that the cautionary effect of uncertainty makes entry up to 70 percent less responsive to tariff reductions on average. In a quantification exercise, the model predicts that if Australia unilaterally reduced tariffs to free trade levels in 2001, the number of traded products would increase by 4 percent. Alternatively, if Australia both reduced tariffs to zero and bound them through WTO commitments, the combined impact of removing caution and the incentive to delay investment would increase the number of traded products by 17 percent. More than half of predicted new product growth is accounted for by reducing uncertainty. In a counterfactual exercise where Australia lowers its tariffs from 1993 to 2001, but does not implement Uruguay Round binding commitments, growth in exporter-product varieties would have been 7 percent lower. These estimates empirically quantify the value of binding tariff commitments for the first time.

In the next section, I describe the Australian trade policy context and semi-parametric evidence for the

[7]In contrast, many countries bound their tariffs at across the board ceilings of 25 or 40 percent, leaving no variation to exploit empirically

role of bindings on trade. In section 3, I develop a *general model* that provides a mechanism for how policy uncertainty affects trade. I then take the predictions of the model to data using reduced form and model-consistent measures of policy uncertainty back to the data in section 4. I quantify the role of uncertainty and conduct several robustness checks. Section 5 concludes.

2 The Application to Trade Policy in Australia

I focus on Australia, a country with a confluence of high quality data and policy variation relevant to uncertainty. In recent history, Australia maintained fairly high applied trade barriers. Unilateral liberalization means there are now large gaps between applied protection and binding commitments. A simple measure of the gap between applied tariffs and bindings is the log of the ratio between the bound and applied MFN tariff rates. The trade policy literature refers to a positive gap in this measure as "binding overhang" or "water in the tariff." In Australia, the gap between applied tariffs in binding ranges from zero to 55 percent on bound tariff lines.

While Australia has low applied tariffs at present, this has not been the case historically. Lloyd's (2008) careful construction of a 100 year time series for Australian tariffs shows that some sectors were highly protected as recently as the early 1990s. There was a legacy of protection for non-competitive industries and political interference in the tariff making process going back to the 1920s (Glezer, 1982). Gradual and, more importantly, *unilateral* liberalization began in the late 1980s and continued into the 1990s.[8] Even in sectors with low applied tariffs, a prospective exporter in the 1990s could look back only a few years justify fear of a high tariff regime.

Historically, policy makers in Australia had adopted a so-called "midway" position in multilateral negotiations. Australia maintained it was neither a developing nor a fully industrialized country and required the flexibility to impose tariffs to protect infant industries with cost disadvantages (Snape et al. 1998). Until the Uruguay Round, only about one fifth of Australia's tariff lines were even bound. And since higher historical tariffs were the starting point for concessions in the Uruguay Round (1986-1994) of multilateral negotiations (see Corden, 1996), Australia's binding commitments today are high and dispersed.

Australia's own Productivity Commission recently noted the prevention of "backsliding" on liberalization as a potential benefit of preferential trade agreements. In their comprehensive review of Australia's trade agreements, the Commission notes that "...even where agreements do not result in a reduction in existing barriers, they can be used to lock in current policies, restricting countries from increasing barriers in the

[8]Coincidentally, journalist Paul Kelly titled his exhaustive book documenting the economic and political upheaval of these reforms "The End of Certainty."

future... in many instances applied tariffs may be low, or even zero, but bound tariff levels might be quite high, and there is a risk that applied tariffs could be increased up to bound levels" (2010, p. 87).

As a case in point, if Australia were to revert all tariffs to their bindings this would substantially shift the tariff profile. In 2001, only 30% of Australia's MFN tariffs are equal to the binding tariff commitment. The magnitude of changes in a reversal to bindings can be large. As the histogram in Figure 1 shows, some MFN tariffs could increase by more than 20 log points in the worst case scenario. Does this degree of binding overhang have real effects on trade or are the low applied tariffs leading to the gaps the only important factor?

A preliminary look at the data suggests the gap between applied and bound rates does affect product entry. In Figure 2(a), the solid line shows a semi-parametric, lowess estimate of the total number of products exported to Australia at the 8 digit tariff line level for 1991-2001 after conditioning on tariff line and year effects. The number of traded products is 20 log points higher for the lowest relative to highest binding gap tariff lines. This is not driven solely by lower tariffs in products with a large gap. After conditioning on applied tariff levels as well, the dashed lowess curve indicates an even stronger negative relationship with respect to the binding gap. Figure 2(b) shows the lowess estimate through the number of traded products and tariffs, again conditional on tariff line and year effects. It is generally downward sloping, as expected. The dashed line shows the same relationship conditional on the binding gap, which is even stronger. Taken together, this evidence indicates a reduction in the extensive margin due to high binding overhang and suggests there is some interaction between applied tariffs and bindings. I develop of model in the next section to understand this mechanism and return the data in Section 4.

3 Model

The basic setup is similar to Chaney (2008) and Helpman et al. (2008), but extended to a deterministic multi-period framework. The world has J exporting countries indexed by j. I consider a single importer, but the model can be extended to a multi-country world. Goods shipped to the importing country are subject to tariffs, which may vary by product and country of origin. I then follow Handley and Limão (2012) by incorporating a stochastic process for tariffs, but extend it to binding tariff commitments. I derive new testable predictions for the role of bindings on entry and their interaction with applied tariffs, which are taken to the data in section 4.

3.1 Preferences

Utility in the importing country is a Cobb-Douglas function over a homogeneous good traded on world markets at zero cost and a continuum of differentiated varieties indexed by i:

$$U = q_0^{1-\mu} \left(\int_{i \in \Omega} q(i)^{\alpha} di \right)^{\mu/\alpha}, \quad \alpha = \frac{\sigma - 1}{\sigma} \tag{1}$$

where $\sigma > 1$ is the elasticity of substitution between varieties and $\mu \in (0,1)$ is the expenditure share of differentiated goods. The total set of varieties available Ω is the union of all domestically produced varieties and those that are imported. Total aggregate income is Y and the utility of a representative consumer is maximized subject to the budget constraint to yields the usual demand function for any particular variety i of

$$q(i) = \mu Y \frac{p(i)^{-\sigma}}{P^{1-\sigma}}. \tag{2}$$

Varieties are differentiated by firm. The price $p(i)$ is the delivered consumer price in the importing country inclusive of tariff and transport costs. The price index is $P = \left[\int_{v \in \Omega} p(i)^{1-\sigma} di\right]^{1/(1-\sigma)}$.

3.2 Production and Tariff Barriers

The homogeneous good is freely traded and produced under CRS such that one unit of the good is produced for $1/w_j$ units of labor in country j. I let q_0 be the numeraire and normalize its price to unity, $p_0 = 1$. Labor market clearing across the numeraire and differentiated goods sectors within a country implies that the wage for country j is w_j. The differentiated goods are subject to ad-valorem tariffs that may vary by exporter j. I let τ_j equal one plus the ad-valorem tariff for goods shipped from country j. Tariffs are paid at the border by consumers on the factory price.

Each firm in country j is identified by its unit labor requirement c_j, which is heterogeneous across firms. The market is monopolistically competitive. Firms chooses prices to maximize profits of $\pi_j = [p_j(c_j)/\tau_j - w_j c_j]q(i)$. This yields the standard markup rule with a delivered price of $p_j(c_j) = \tau_j w_j c_j/\alpha$. Combining the the markup rule, consumer demand and variable costs, the per period operating profits of an exporter in country j are

$$\pi_j(c_j) = A_j \tau_j^{-\sigma} c_j^{1-\sigma} \tag{3}$$

where $A_j = (1-\alpha)\mu Y \left[\frac{w_j}{P\alpha}\right]^{1-\sigma}$ summarizes exporter cost and importer demand conditions.

I assume there is a distribution of unit costs, $G(c)$, that summarizes the heterogeneity in productivity $(1/c)$ within each country and is bounded below at c^L. I index variation in the lower bound across exporting countries by M_j. Then the lowest unit cost firm in country j is $c_j^L = M_j c^L$.

3.3 Entry, Exit and Sunk Costs

There is a fixed cost of market entry K_e paid by a firm to begin exporting. Entry costs cover the expenses of setting up a distribution network, marketing, on-site visits or agency costs, tailoring products to local markets, and complying with safety regulations. There are no fixed entry or per period maintenance costs in a firm's domestic market. Since operating profits are always positive, albeit potentially quite small, every firm sells in its home market. A subset of firms pay the entry cost and begin exporting if their unit costs are below a threshold cutoff level. Following Melitz (2003), exit from exporting is induced by an exogenous death shock with probability δ. A firm that is hit by the death shock exits immediately without recouping its sunk costs.

In a deterministic environment, where $\pi_j(t) = \pi_j$ in the future, the firm will enter an export market if the net present discounted value of entry is positive, such that

$$V^D = \frac{\pi_j}{1-\beta} - K_e \geq 0 \tag{4}$$

The superscript D denotes a "deterministic" tariff regime. The discount factor combines the true discount rate ρ and the death shock such that $\beta = (1-\delta)/(1+\rho)$. Free entry implies that in equilibrium $V^D = 0$ for the marginal entrant. Imposing this condition yields a multi-period zero cutoff profit threshold for unit labor costs c_j^D

$$c_j^D = \left[\frac{A_j \tau_j^{-\sigma}}{(1-\beta)K_e}\right]^{1/(\sigma-1)} \tag{5}$$

All firms with unit costs below c_j^D will pay the entry cost and begin exporting. It is straightforward to derive that the elasticity of c_j^D to a once-and-for-all change in τ is $\frac{\sigma}{\sigma-1}$.[9]

3.4 A stochastic framework for trade policy uncertainty

In practice, the level of future tariffs is uncertain. Many factors can affect the formation of trade policy over time. I take shocks to trade policy as given and do not explicitly model their source. Tariffs are a random

[9] This is higher than the usual elasticity of unity because tariffs are paid at the border, rather than as part of the firm's variable trade cost technology.

variable with two sources of variation: uncertainty over the timing of policy changes, and uncertainty over the magnitude of those changes when they arrive. Even though the outcome of policy changes is unknown *ex-ante*, firms can form expectations over the likely tariff outcomes.

To model tariff uncertainty, I use the underlying framework in Handley and Limão (2012) and I assume shocks to the path of tariffs arrive with probability γ per unit of time.[10] When a shock arrives, a policy maker sets a new tariff τ'. Firms know the value of γ and can assign probability measures to different tariff outcomes. The space of potential tariff outcomes and their likelihood are summarized by the distribution function $H(\tau')$ with support $[1, \tau_{max}]$. The support spans from free trade ($\tau = 1$) to autarky when letting $\tau_{max} \to \infty$.

3.5 Entry and Exit under Uncertainty

Under a stochastic tariff process, there is an option value of waiting with a structure similar to Baldwin and Krugman (1989). While the current tariff is known, future profit flows are subject to the stochastic process for tariffs. The firm's decision to enter an export market is an optimal stopping problem. Firms can be divided into exporters, state 1, and non-exporters, state 0. The value of being an exporter in the current period is V^1. A firm that is in state 1 exits only when hit by the death shock. Non-exporters hold an option value of waiting to enter in the future V^0. Non-exporters will enter a foreign market only when the value of exporting less sunk entry costs exceeds the value of waiting such that $V^1 - K_e \geq V^0$.

The decision rule for each firm is defined by the trigger tariff τ_1 that makes the firm indifferent between entry and waiting. For each firm, identified by its unit labor requirement c, the entry trigger τ_1 implicitly solves the indifference condition

$$V^1(\tau_1) - K_e = V^0(\tau_1) \tag{6}$$

A firm will enter the export market if $\tau_t \leq \tau_1$.

The solutions for the current values of entry and waiting are derived in the appendix. The entry margin corresponds to the firm with unit labor requirement c^U that is indifferent to entry or waiting at time t.[11] For this marginal firm, the current tariff equals the entry trigger such that $\tau_t = \tau_1$. The expression in (6) defines a zero cutoff profit condition for the entry margin. I solve directly for c^U and express it in terms of

[10]In continuous time, a similar Poisson process for the arrival of tax policy changes can be found in Rodrik (1991) and Hassett and Metcalf (1999).

[11]Superscript U denotes the "uncertain" environment in contrast the the "deterministic" environment D.

an uncertainty component $\Theta(\tau_t)$ and the deterministic cutoff c^D:

$$c^U = \Theta(\tau_t) \times c^D \tag{7}$$

$$\text{where } \Theta(\tau_t) = \left[\frac{1-\beta+\beta\gamma\Delta\left(\tau_t\right)}{1-\beta+\beta\gamma}\right]^{\frac{1}{\sigma-1}} \tag{8}$$

As shown in the appendix, $\Theta(\tau_t) \leq 1$ since $\Delta(\tau_t) \leq 1$. For a given current tariff, uncertainty over the tariff generates a lower cost cutoff than a deterministic model. The productivity premium necessary to overcome this hurdle is the ratio of c^D and c^U, or $\frac{1}{\Theta}$. T he model includes the deterministic environment as a special case. When $\gamma = 0$, the option value of waiting vanishes. The zero cutoff profit threshold collapses to the deterministic expression in (5).

The expression for $\Delta(\tau_t)$ captures the random variation in the tariff conditional on a policy shock arrival. In the appendix, I show the following:

$$\Delta\left(\tau_t\right) - 1 = \left(1 - H(\tau_t)\right)\left[\frac{E(\tau^{-\sigma} \mid \tau \geq \tau_t) - \tau_t^{-\sigma}}{\tau_t^{-\sigma}}\right] \leq 0 \tag{9}$$

I interpret $\Delta\left(\tau_t\right) - 1$ as the expected proportional reduction in operating profits that occurs following a bad shock. The leading term $(1 - H(\tau_t))$ is the probability of a shock that exceeds the current tariff for the marginal firm. The term in brackets is the expected proportional loss in profits, starting from τ_t, if a bad shock arrives. The inequality is always strict except when the current tariff is at the maximum of the tariff distribution, in which case $c^D = c^U$.

In Handley and Limão (2012) the focus is on how PTAs reduce the arrival rate of bad shocks, primarily through changes in γ. I focus instead on how bindings reduce policy uncertainty via their effect on the $H(\cdot)$ distribution. A credible WTO binding is the maximum tariff permitted by WTO rules. The commitment to bind tariffs is a constraint on observable tariff outcomes such that the distribution of future tariffs, $H(\tau')$, is censored at the binding. By analogy to Tobit regression, censoring captures the idea that a policy maker might want to set a tariff above the binding but WTO legal constraints mean that only the binding tariff is actually observed.

I let B denote the level of the binding, which must be below the maximum of the unbound tariff distribution to be effective, $B < \tau_{max}$. Binding commitments induce a mixed discrete and continuous distribution over tariffs. A formal statement of the *bound* tariff distribution appears in the appendix. When a policy shock arrives, the new tariff is a random draw from $H(\tau')$. There is a discrete probability $H(B)$ that the tariff draw is below the binding. But with probability $1 - H(B)$, the tariff draw is above the binding and only the bound tariff rate is observed.

It is straightforward to show that binding the tariff distribution will increase the cutoff. The binding augmented version of the profit loss term in equation (9) is[12]

$$\begin{aligned} \Delta(\tau_t, B) - 1 &= \frac{(1 - H(B))[B^{-\sigma} - \tau_t^{-\sigma}] + [H(B) - H(\tau_t)]E(\tau^{-\sigma} \mid \tau_t < \tau < B)}{\tau_t^{-\sigma}} \\ &= \frac{(1 - H(B))[B^{-\sigma} - E(\tau^{-\sigma} \mid \tau > B)]}{\tau_t^{-\sigma}} + [\Delta(\tau_t) - 1]. \end{aligned} \tag{10}$$

To understand how binding effects the cutoff, consider the two limits of the bound expression in the first line above. When $B > \tau_{max}$ so that so that $H(B) = 1$, then (10) is equivalent to (9). Alternatively, when the binding is decreased to $B = \tau_t$, the tariff is at the max of the bound distribution. The profit loss expression in (10) goes to zero, replicating the deterministic cutoff. Turning to the second line, the first term captures the shift in probability mass from the upper tail of the tariff distribution to the binding $\Delta(\tau_t, B) - \Delta(\tau_t) > 0$. This limits losses in the worst case and increases the cutoff. In effect, making a credible commitment not to increase tariffs in the future can move firms toward the deterministic cutoff.

This is an example of the "bad news" principle identified by Bernanke (1983). Even though another policy shock could induce a new tariff that is lower than the current tariff, it is only the prospect of a bad shock that affects the decision of whether to enter today. It holds despite the convexity of profits in tariffs. When a firm enters, it weighs the expected PDV of profits from entering today against the value of waiting for a better shock in the future. Because good news in the future is offset by the opportunity cost of entry, only bad news matters when the entry investment is irreversible.

3.6 Implications of Uncertainty for the Entry Cutoff

Uncertainty about future trade policies delays entry at the margin relative to the deterministic model. Reducing uncertainty will lead prospective firms to bring entry forward even if applied tariffs remain unchanged. Uncertainty also makes firms on the margin more cautious. For a given tariff reduction, the elasticity of the entry cutoff to changes in tariffs is attenuated by uncertainty. These results, delay and caution, can be derived analytically and have important implications for policy.[13] Detailed derivations appear in the appendix.

Trade policy uncertainty generates first order reductions in the entry margin. These are summarized in the following proposition.

[12] A derivation of $\Delta(\tau_t, B)$ is in the appendix.

[13] These caution and delay effects are related to similar findings for investment in studies of firm-level response to demand shocks under uncertainty (e.g. Bloom et al., 2007).

PROPOSITION 1 ***[Delay]*** *Higher bindings reduce the entry cutoff, c^U, by delaying investment in market entry, and making entry more sensitive to the policy shock arrival rate. In elasticity terms:*

(a) The cutoff is decreasing in the binding

$$\varepsilon(B) = \frac{d \ln c_t^U}{d \ln B} = -\frac{\sigma}{\sigma - 1} \left(\frac{\beta\gamma}{(1 - \beta + \beta\gamma\Delta(\tau_t, B))} \left(\frac{(1 - H(B))B^{-\sigma}}{\tau_t^{-\sigma}} \right) \right) < 0$$

(b) The cutoff is decreasing arrival rates

$$\varepsilon(\gamma) = \frac{d \ln c_t^U}{d \ln \gamma} = \frac{\beta\gamma}{\sigma - 1} \left[\frac{1 - \beta}{(1 - \beta + \beta\gamma)[1 - \beta + \beta\gamma\Delta(\tau_t, B)]} \right] [\Delta(\tau_t, B) - 1] \leq 0$$

(c) The cutoff is more responsive to arrival rates when bindings increase $\frac{\partial\varepsilon(\gamma)}{\partial B} = \frac{d^2 \ln c_t^U}{d \ln \gamma \cdot dB} \leq 0$

PROOF:(see appendix)

This proposition isolates the effects of policy shock timing and magnitudes into two components. First, binding reductions can increase entry, even if they do not constrain the current applied tariff, by lowering tariffs in the worst case scenario and bringing entry forward. In an environment where policy shocks cannot be eliminated, lower bindings can raise trade even if the binding is above the current period applied tariff. Second, increases in the arrival rate of policy shocks reduce entry. In the deterministic limit $\varepsilon(\gamma)|_{\gamma=0} = 0$ and this delay effect vanishes.[14] Finally, part (c) shows that lower bindings can reduce or even eliminate the effect of higher arrival rates on the cutoff. For example, when there is no binding gap and $B = \tau_t$, then $\varepsilon(\gamma) = 0$.

These results follow from the option value of waiting. Future tariffs could have a lower expected value than current tariffs and some firms would still delay entry. Even if a more favorable tariff regime is on the horizon, delaying entry may be optimal. Similarly, when current tariffs are low and expected tariffs are high, firms on the margin will wait to enter.

PROPOSITION 2 ***[Caution]*** *The entry cutoff c^U is less elastic with respect to a given tariff change under uncertainty ($\gamma > 0$).*

$$\begin{aligned}
\varepsilon^U(\tau_t) &= \frac{d \ln c_t^D}{d \ln \tau_t} + \frac{d \ln \Theta_t}{d \ln \tau_t} \\
&= -\frac{\sigma}{\sigma - 1} \left[1 - \left(\beta\gamma \frac{(1 - H(B))B^{-\sigma} + (H(B) - H(\tau_t))E[\tau^{-\sigma} \mid \tau \in (\tau_t, B)]}{[1 - \beta + \beta\gamma\Delta(\tau_t, B)]\tau_t^{-\sigma}} \right) \right].
\end{aligned}$$

[14] This same insight appears in Handley and Limão (2012), but part (c) does not.

PROOF:(see appendix)

The leading term above is the deterministic elasticity. It is attenuated by the bracketed term, which is less than or equal to one. In absolute magnitudes $\mid \varepsilon^U(\tau_t) \mid < \mid \varepsilon^D(\tau_t) \mid$ and the responsiveness of the entry margin is reduced under uncertainty. The two exceptions (limiting cases) are when $\gamma = 0$ (i.e. tariffs are deterministic) or when τ_t is already at the maximum of the tariff distribution.[15]

On the extensive margin, binding reductions could be just as effective as applied tariff reductions for increasing trade. The figurative "insurance" against backsliding through binding commitments would be relevant if prospective entrants place some probability weight on the possibility of large scale tariff reversals. In theory, further reductions in binding commitments through WTO negotiations would be meaningful. This is a testable implication of the model. Consider a current tariff τ_0 that is below its binding B_0. Suppose the current tariff and binding are changed by $d\ln\tau$ and $d\ln B$, respectively. Using Propositions 1 and 2, the comparative static for the *net* change in the entry cutoff $d\ln c^U$ is computed as follows

$$\begin{aligned} d\ln c^U &= \varepsilon^U(\tau_0)d\ln\tau + \varepsilon(B_0)d\ln B \\ &= \varepsilon^D(\tau_0)d\ln\tau - \varepsilon(B_0)\times(d\ln B - d\ln\tau) + r_0 \times d\ln\tau \qquad (11) \\ \text{where } r_0 &= \frac{\sigma}{\sigma-1}\left(\frac{\beta\gamma}{(1-\beta+\beta\gamma\Delta)}\left(\frac{(H(B_0)-H(\tau_0))E[\tau^{-\sigma} \mid \tau\in(\tau_0,B_0)]}{\tau_0^{-\sigma}}\right)\right). \end{aligned}$$

The first term is the deterministic elasticity. The second term captures relationship between simultaneous binding and tariff changes. If the binding is unchanged, say in a unilateral tariff reduction, then the impact on the entry cutoff is reduced. When both the binding and tariff change by the same amount, $d\ln B - d\ln\tau = 0$, the second term drops out. The third term is the residual uncertainty about tariff outcomes in the policy space between τ_0 and the binding B_0. Residual uncertainty will reduce the elasticity of the cutoff if the gap between B_0 and τ_0 is large and the probability mass in that range of the tariff distribution is high. This comparative static result is summarized in the following corollary to propositions 1 and 2.

COROLLARY 1 ***[Bound tariff changes]*** *Tariff changes accompanied by equal or greater changes in binding commitments will generate more new entry than unbound, unilateral tariff changes.*

When tariffs are reduced unilaterally, without constraining future policy makers through binding, the impact on the entry cutoff is mitigated. I confirm the broader implications of this prediction in the quantification and counterfactual exercises.

[15] A related results without bindings appears in Handley and Limão, (2012)

4 Empirical Evidence

The empirical strategy is meant to test if exporters place positive probability weight on a reversal to binding tariff levels. First, I test the reduced form predictions of the model. Second, I use structural measures of uncertainty derived from the model to investigate the mechanism in more detail and quantify the role of policy uncertainty. The panel data approach exploits the cross-section and time variation in tariffs and bindings to identify the probability that a product is traded.

4.1 Estimation Method

I estimate the model on disaggregated product data. Data on firms from a multitude of potential import partners are not available. A reasonable measure of firm entry is whether a disaggregated product is traded.[16] Even if firm data were available, it would be difficult to identify the set of potential exporters and estimate entry probabilities at the tariff line level for the universe of all firms. A method for evaluation of trade policy reforms under uncertainty with more widely available product data is a contribution of this paper.

I extend the model to multiple industries, by allowing tariffs and bindings to vary by products. Using v to denote a product, trade is observed if the unit cost of the most productive firm in country j is below the cutoff for a particular product, specifically $c_j^L < c_{tjv}^U$.

The unit cost of the marginal firm c_{tjv} from country j in product v is not observed, but it must equal the cutoff threshold $c_{tjv} = c_{tj}^U$. The ratio of the cutoff for the marginal firm in product v to that of the most productive firm in the industry c_j^L can be defined in terms of observables as a latent variable.[17] If the expected PDV of entering today is greater than or equal to the fixed cost of entry, I observe the decision of at least one firm to enter when a product is traded. I define a latent variable Z_{tjv} for each exporter(j)-product(v)

$$Z_{tjv} = \left(\frac{c_{tj}^U}{c_j^L}\right)^{\sigma-1} = \frac{\Theta_t^{\sigma-1}\tau_{tjv}^{-\sigma}A_{tj}(c_{jI}^L)^{1-\sigma}}{(1-\beta)K_{tjv}}$$

where the second equality follows from substitution of equations (7) and (5) for c_{tj}^U. If $Z_{tjv} \geq 1$ for at least one exporter in country j, then trade is observed in that exporter-product pair.

Taking logs and substituting for Θ using equation (8) yields

$$z_{tjv} \propto -\sigma \ln \tau_{tjv} + \ln\left[\frac{1-\beta+\beta\gamma\Delta}{1-\beta+\beta\gamma}\right] + d_{jt} + d_{jv}. \tag{12}$$

[16]The evidence of firm level entry following trade liberalizations is confirmed in disaggregated product level studies such as Kehoe and Ruhl (2013) and Debaere and Mostashari (2010).

[17]A similar cross-country latent variable formulation is used in Helpman et al. (2008). Armenter and Koren (2010) and Santos Silva et al. (forthcoming) have also used product or sector level trade indicators to proxy for the firm extensive margin.

I assume that sunk export costs vary across each exporter and products such that $\ln K_e = k_{jv}$. A set of exporter-year fixed effects d_{jt} and exporter-product effects d_{jv}, control for unobserved variables. The exporter-year effect $d_{tj} = (1-\sigma)\ln M_{tj} + (1-\sigma)\ln w_j + y_t + (\sigma-1)p_v$ encompasses unobserved heterogeneity in aggregate productivity, exporter wages, and demand conditions. The exporter-product effect $d_{jv} = k_{jv} + \ln \mu_v + (1-\sigma)\ln c^L_{jv}$ combines unobserved heterogeneity in entry costs, expenditure shares, and minimum unit costs. Trade is observed when $z_{tjv} = \ln(Z_{tjv})$ is positive.

4.2 Data Implementation and Sample

A description of the data sources appears in the appendix. Here I focus on construction of the regression samples, tariffs and uncertainty measures.

Applied tariffs and bindings are set at the 8 digit level of detail on Australian imports. Applied tariff data are obtained from UNCTAD TRAINS database via the World Integrated Trade Solution (WITS). Tariff data are not available for 1992, 1994, or 1995 via WITS or any other source. A large number of developing country exporters are eligible for preferences under one or more programs. TRAINS also contains data on all preferential tariff schedules and country eligibility. However, these preferential tariffs require additional documentation and compliance costs and are not always utilized (Pomfret et al., 2010). I construct tariff line measures of the ad valorem applied tariff (i.e. 1+ tariff rate) for both the MFN tariff and the preferential tariff, if any, for each country-tariff line combination.

The data on bindings at 8 digit level for Australia's Uruguay Round commitments come from the WTO's consolidated tariff schedules (CTS) accessible through the Tariff Analysis On-line database. These bindings were effective in 1996 and certified by the WTO in 2001 the final year of the sample. Australia had made previous commitments on a smaller set of products during the Tokyo Round. Data on these bindings at the 8 digit level were obtained from the WTO Integrated Database Version 1.0 on CD-ROM.[18]

Other forms of protection are limited. Australia abolished quotas and removed most other quantitative import restrictions in a process known as "tariffication" leading up the Uruguay Round (Snape et al. 1998). This means that measured trade barriers are homogeneous across products as the primary instrument of protection is an ad-valorem tariff rate. Moreover, the effective rate of domestic industry assistance has been constant at about 5% since the 1990s (Productivity Commission, 2010, p. 48).

I obtained annual import data by value for duty at the 10-digit level of detail for Australia from 1990 to 2001. These product classifications are extremely detailed. For instance, Australian Customs tracks 67 different varieties of tubes and pipes. If I break the data down to its nuts and bolts, literally, I find there are ten different varieties of bolts, which can be fastened with two types of nuts. For consistency with the

[18]Further details on the Tokyo Round bindings data are in the data appendix.

tariff data, I aggregate to the 8 digit level. The reporting threshold for customs valuation also varies over time and I imposed a flat threshold across all years of $10,000 AUD as the cutoff for whether a product was traded.

There are significant changes to the Harmonized System of product classification in 1996 as well as smaller modifications to 8 digit codes annually by the Australian Bureau of Statistics.[19] In order to make consistent comparisons across time, I concorded 8 digit codes using the method of Pierce and Schott (2009). I then match the import and tariff data at the concorded 8 digit level to construct a country-product by year panel of all traded and non-traded products and their associated trade policies. Only GATT/WTO member countries are included in the set of exporters.

The final sample contains 3,770,862 exporter-product-year observations for 1991, 1993, and 1996-2001. Table 1 reports summary statistics. Within the sample, the average applied tariffs are low at approximately 4.5 log points. The average binding is nearly twice as high at 9.4 log points. Many exporters are granted preferences, but due to preference erosion the average margin of preference in the data is only 0.4 log points. Nevertheless, the regressions that follow control for several aspects of preferential access.

The baseline estimation is based on an unbalanced panel. As noted above, many products were unbound prior to 1996 and so no measure of bindings is available. In addition, the membership of the GATT/WTO changes over time so new countries enter the sample in later years. Regressions on a balanced panel using only the subsample of products bound in the Tokyo Round or only using the post Uruguay Round sample (1996-2001) are in the appendix.

4.3 Reduced form model estimation

The latent variable equation in (12) differs from a deterministic model due to the bracketed term, which is non-linear in the parameters of interest. Using the elasticities from propositions 1 and 2, a reduced form estimation equation can be obtained. The model predicts the the cutoff is decreasing in the tariff and the gap between bindings and applied tariff according to equation (11). The estimating equation to test these relationships is

$$z_{tjv} = a_\tau \ln \tau_{tjv} + a_B \ln \left(\frac{B_{tv}}{\tau_{tjv}} \right) + d_{jt} + d_{jv} + \varepsilon_{tjv} \tag{13}$$

where ε_{tjv} is an i.i.d error term. The parameter $a_\tau = (\sigma - 1)\varepsilon(\tau) = -\sigma < 0$, which is the rescaled deterministic tariff elasticity of the cutoff.[20] The parameter $a_B = (\sigma - 1)\varepsilon(B) < 0$ and measures how changes in the gap

[19] I used a correspondence file for all code changes from 1988 to 2001 obtained directly from the Australian Bureau of Statistics.
[20] Here I assume the residual uncertainty term, r_0, is small. Otherwise it is estimated as part of the coefficient on the tariff.

between applied tariffs and bindings effects entry. If there is zero probability of future tariffs at or above the bound rate, then $a_B = 0$.

Let T_{tjv} be a binary indicator defined as $T_{tjv} = \mathbf{1}[z_{tjv} > 0]$. I model the probability that a product is traded as $p^{(T=1)}_{tjv} = \Pr.(T_{tjv} = 1 \mid Xb) = F(Xa)$ where $F(\cdot)$ is a CDF. The estimating equation is

$$p^{(T=1)}_{tjv} = F\left[a^*_B \ln\left(\frac{B_{tv}}{\tau_{tjv}}\right) + a^*_\tau \ln \tau_{tjv} + d_{jt} + d_{jv}\right]. \tag{14}$$

The high dimensional set of exporter-year and exporter-product fixed effects mean that a standard Probit model would be inconsistent and not computationally feasible. I assume instead that $F(\cdot)$ is linear and estimate a linear probability model (LPM) using OLS. I conduct robustness checks of the specification itself and the estimation method in section 4.6.1.

One final issue of applied tariff measurement must be handled. Some exporters receive tariff line preferences for developing countries (with the exception of New Zealand, which has a PTA). Data on preference utilization is not available and for these tariff lines the true applied tariff is unobserved. To handle this, I construct a dummy variable $D_{pref} = 1$ for countries where the preferential country-product specific preference margin is positive, $\ln \tau_{tv,MFN} - \ln \tau_{tjv} > 0$. The unobserved applied tariff is $\tau_{tjv} = (1 - D_{pref}) \times \ln \tau_{tv,MFN} + D_{pref} \times \ln \tau_{tjv}$ and this can be rearranged to yield $\tau_{tjv} = \tau_{tv,MFN} - D_{pref} \times (\ln \tau_{tv,MFN} - \ln \tau_{tjv})$. Substituting this expression into (14) for τ_{tjv}

$$p^{(T=1)}_{tjv} = F\left[a^*_B \ln\left(\frac{B_{tv}}{\tau_{tjv}}\right) + a^*_\tau \ln \tau_{tv,MFN} + a^*_{pref}(\ln \tau_{tv,MFN} - \ln \tau_{tjv}) + d_{jt} + d_{jv}\right]. \tag{15}$$

Then a^*_τ identifies the elasticity of entry with respect to the MFN tariff and a^*_{pref} identifies an adjustment factor when there is a positive preference margin $(\ln \tau_{tv,MFN} - \ln \tau_{tjv})$.

The advantage of the reduced form approach is that it requires a minimal set of identifying assumptions on parameters and the tariff distribution $H(\tau')$. The starred parameters (a^*_τ and a^*_B) in a LPM are scaled into the marginal effects on the probability a product is traded. The elasticity of substitution σ can vary across industries, in which case a^*_τ identifies the average elasticity scaled in its marginal effect. I also require that exporters within a tariff line defined product form the same expectations, using the same tariff distribution, about future policies. This is necessary to identify the average elasticity of the binding gap conditional on the current trade policy. The assumption is consistent with a rational expectations environment where there are no arbitrage opportunities.

Results for the full panel are in Table 2. I find negative and significant effects of the both the tariff and the log of the binding gap. In column (2), I add the preference margin control, which is positive and significant. Including this control is important, as it increases the magnitude of the tariff marginal effect but it has little effect on the binding gap coefficient. When the binding gap is omitted from the regression, as in column (3), the marginal effect of the tariff is much lower. This precisely what classic omitted variable bias would predict and the model tells us why. In the corollary to propositions 1 and 2, the applied tariff elasticity is attenuated by the binding gap.

4.4 Structural-based model estimation

I now return to (12) and derive a model-consistent structural measure of trade policy uncertainty with bindings. In the deterministic limit where $\gamma = 0$ the bracketed uncertainty term in (12) drops out entirely. Since I ultimately test for the presence of uncertainty, I take $\gamma = 0$ as a testable null hypothesis and linearize around this point. The first-order Taylor approximation to $\ln \Theta_{tjv}^{\sigma-1}$ around $\gamma = 0$ is

$$\left.\frac{d \ln \Theta_{tjv}^{\sigma-1}}{d\gamma}\right|_{\gamma=0} \approx \frac{\beta\gamma}{1-\beta}(\Delta(\tau_{tjv}) - 1). \tag{16}$$

The linearized uncertainty term parsimoniously represents the two components of the uncertainty process: the magnitude of the expected proportional loss in profits given a policy shock arrives is captured by $\Delta(\tau_{tjv}) - 1$; the arrival rate of trade policy shocks appears linearly in γ. Estimation requires measures of the profit losses that could occur in a reversal.

A strength of the analytical simplicity of this model and the focus on trade policy is that measures of the expected profit loss can be constructed from tariff data. I discretize the expected loss for a reversal to the binding tariff from the applied MFN tariff with probability $p_B = 1 - H(\tau_{tjv})$. The discrete decomposition is

$$\Delta(\tau_{tjv}) - 1 = -p_B \frac{\tau_{tjv}^{-\sigma} - B_{tv}^{-\sigma}}{\tau_{tjv}^{-\sigma}} = -p_B U_{tv}^B \tag{17}$$

The uncertainty measure, U_{tjv}^B, is bounded below at zero and bounded above at 1 for a reversal to total autarky.

For all exporters, I set the applied tariff as the MFN tariff such that $\tau_{tjv} = \tau_{tv,MFN}$. For any tariff line where the bound tariff is above the MFN tariff, the "binding uncertainty" measure U_{tjv}^B is positive. I assume that $\sigma = 4$ for the elasticity of substitution in my baseline estimates, but show these are robust to the choice of σ.[21] The measure provides a way to estimate the impact of uncertainty through the lens of the model. For example, "Windscreens of toughened (tempered) safety glass of a kind used as components in passenger motor vehicles" had an MFN tariff of 15% and bound rate of 37.5% in 1993. This corresponds to a profit

[21] Bernard et al. (2003) estimate that $\sigma = 3.8$ using U.S. firm level trade data.

loss measure, U^B, of 51%, when $\sigma = 4$. After the Uruguay Round, the bound rate was reduced to 19% and by 2001 the MFN tariff was 10%, for a correspondong potential profit loss of 27%. Disentangling the role of the tariff reduction and the binding reduction are the focus of the following estimation strategy.

Substituting the uncertainty measure into equation (12) yields

$$z_{tjv} = -\sigma \ln \tau_{tjv} - p_B \gamma \frac{\beta}{1-\beta} U^B_{tjv} + d_{jt} + d_{jv} + \varepsilon_{tjv} \tag{18}$$

The estimating equation using the first-order approximation for the uncertainty measure and the preference margin variable specified above is

$$p^{(T=1)}_{tjv} = F[b^*_B U^B_{tv} + b^*_\tau \ln \tau_{tv,MFN} + b^*_{pref}(\ln \tau_{tv,MFN} - \ln \tau_{tjv}) + d_{jt} + d_{jv}]. \tag{19}$$

As in the reduced-form regressions, there is no variation of the binding or MFN tariff within exporter-product, so standard errors are clustered at the 8 digit tariff line by year level.

The starred parameters (b^*_τ and b^*_B) are scaled into the marginal effects on the probability a product is traded. The estimated, unstarred coefficients can still be interpreted in the context of the model. The elasticity of sales to applied tariffs is negative and estimated by the parameter $b_\tau = -\sigma$ up to a scale factor. The negative impact of uncertainty is estimated up to scale by the parameter $b_B = \frac{\beta}{1-\beta}\gamma \cdot p_B$ where the term in discount factors is a positive constant. These coefficients are proportional to the probability weight placed on reversals to the binding, given by $\gamma \cdot p_B$. Negative coefficients indicate exporters in the average tariff line place some weight on bad news when making entry decisions.

Before reporting the baseline results, I discuss the sources of identifying variation and potential bias. Exporter-product effects control for important time-invariant unobserved heterogeneity influencing the entry decision in the model: sunk costs, expenditure shares, and the unit cost minimum. They also control for endogenous protection and other exporter-product specific trade barriers that might confound results. The exporter-year effects handle aggregate wages, exchange rates, the price index, and demand conditions. The size and direction of feedback from tariff and binding reductions to these sources of aggregate time-variation is unclear. Their omission could overstate the effect of tariffs and uncertainty if tariff reductions are correlated with favorable shocks in Australia and its trade partners, e.g. higher incomes, better terms of trade. Alternatively, Australian trade liberalization may bid up wages in its export partners, increasing their costs.[22] Likewise, liberalization could make the Australian market more competitive for all firms via price index effects. Such feedback could attenuate the effect of tariff and binding reductions on entry or

[22] Wages are fixed endogenously in the theory, but this channel operates in other GE heterogeneous firms models.

trade flows.

Results from the baseline structural measure of binding uncertainty are in the first 3 columns of Table 3. Column (1) includes no time effects. Both tariffs and the uncertainty measure are negative and significant. Magnitudes are large, most likely because tariff reductions are negatively correlated with increases in aggregate income in both Australia and its trade partners. When controlling only for year effects in column (2), magnitudes fall by about one third, but remain significant. Finally, in column (3) I add the full set of exporter-year effects suggested by the model. This only effects the preference margin control; it's coefficient falls by about 25% as there are many exporter-specific requirements to qualify for preferential rates. Column (3) is the preferred specification that I will use in the quantification exercises that follow. It is important, given this discussion, that all inference should be interpreted conditional on the set of exporter-product and exporter-year effects.

In order to give the coefficients a structural interpretation, I check whether interactions between MFN tariffs and the preference margin measure are driving the negative coefficient on binding uncertainty. For example, exporters with preferential market access may be less sensitive to binding uncertainty. To investigate this possibility I run three additional checks. First, I omit the preference margin in column (4). This biases the applied MFN tariff coefficient downward as in the reduced form regressions, but has almost no effect binding uncertainty. Second, I drop all exporter-product observations that have a positive preference margin in column (5). This increases the magnitudes of both the tariff and binding uncertainty coefficients, but not substantially. Third, it's possible exporters with preferential market access feel more or less secure, even if preference erosion means that the preferential tariff equals the MFN tariff. In column (6) I drop all observations with preferential market access even if the preference margin is zero, which is about 80% of the sample. This is does not substantially change the coefficients either. The ratio of the binding uncertainty to tariff marginal effect is 0.21 to 0.22 regardless of how preferences are handled.

4.5 Quantification

4.5.1 Elasticities

The first-order approximation used to compute the uncertainty terms decouples the elasticity on applied tariffs from the uncertainty measure. In order to measure and test the cautionary effects derived in Proposition 2 and elaborated in the Corollary, I need to account for the fact that the uncertainty measure is a function of tariff and binding levels. The elasticity of entry to tariff and binding changes is computed by log differentiation of the uncertainty measure. In terms of the model, the estimated elasticity of product entry to tariff reductions, $e(\tau)$ is the sum of the direct effect to current profits, the second term in (19), and the change to future profits if a reversal occurs, the first term in (19):

$$\begin{aligned} e(\tau) &= b_\tau + b_B \frac{\partial U^B_{tv}}{\partial \tau} \tau_{tv,MFN} \\ &= b_\tau - b_B \sigma \times \frac{B^{-\sigma}_{tv}}{\tau^{-\sigma}_{tv,MFN}} \\ &= -\sigma \left[1 - \gamma p_b \frac{\beta}{1-\beta} (1 - U^B_{tv}) \right]. \end{aligned} \tag{20}$$

This is simply the first order approximation to the cautionary effect derived in Proposition 2.[23] The elasticity of entry to applied tariff changes will depend on the probability of reversals, $\gamma \cdot p_B$, and their proportional magnitudes $B^{-\sigma}/\tau^{-\sigma} = 1 - U^B$.

Proposition 1 shows that the elasticity of entry is reduced by increases in bindings. I can also use the empirical model structure to obtain the elasticity of entry to changes in binding levels following the same computation as above:

$$\begin{aligned} e(B) = b_B \frac{\partial U^B_{tv}}{\partial B} B_{tv} &= -b_B \sigma \times \frac{B^{-\sigma}_{tv}}{\tau^{-\sigma}_{tv,MFN}} \\ &= -\sigma \gamma p_b \frac{\beta}{1-\beta} (1 - U^B_{tv}) < 0. \end{aligned} \tag{21}$$

The elasticity of entry to binding changes and the cautionary effect from above are symmetric. If bindings are reduced by the same percentage as applied tariffs, the cautionary effect is exactly offset "as if" tariffs were deterministic. This result follows from the Corollary to Propositions 1 and 2.

These elasticity quantifications depend on the parametric assumptions on σ used for constructing the uncertainty measure. To assess the robustness of the assumption and the following quantification, I run a set of robustness checks using estimated measures of σ that vary across industries from Kee et al. (2009). These estimates are noisy and missing at the 6 digit level in many cases. As such, I take the median elasticity within each 2 digit Chapter (HS2) and the 21 sections of the HS and use them to compute the binding uncertainty measure.

Estimates appear in Table 4 next to the baseline parametric results when $\sigma = 4$. Two features stand out. First, there is little effect on the tariff marginal effect or the preference margin across the alternative measures of σ. Second, the marginal effect on the binding uncertainty measure increases in magnitude. To compare the impact of reducing bindings versus unilaterally reducing applied tariffs, I turn to the lower panel of Table 4. Caution and delay effects are large and evident after I compute the elasticities at the mean

[23] In the approximation, it is not assured that the term in brackets is less than one, as in the theoretical model, since $\frac{\beta}{\beta-1} > 1$ whenever $\beta > 0.5$. Nevertheless, the estimates from the econometric model below do satisfy this restriction in practice.

of the uncertainty measure using expressions (20) and (21). The elasticity of entry to tariff reductions is reduced by 55 to 88 percent due to the caution effect across all 3 parameterizations of σ. Delay effects are also important. The elasticity of the probability of being traded increases by 9 to 12 percent for every 1 percent decrease in bindings.

Using estimated σ's introduces potential sources of bias and inconsistency. Given the theoretical predictions of the model, elasticities of substitition estimated without accounting for applied policy uncertainty could be biased, as the reduced form results indicated. Since the baseline parametric elasticities when $\sigma = 4$ lie between the HS2 and section level estimates, I use them in the quantification exercise that follows.

As a final specification check, I note that the tariff marginal effect is an unconstrained measure of σ scaled into its marginal effect. The regressions estimate of b^*_τ is an average of σ scaled into its marginal effect across all tariff lines whereas this coefficient is constrained in the uncertainty measure. When σ is allowed to vary across sectors in columns (2) and (3) this does not change the tariff coefficient, which suggests any potential bias is minimal. As a further check whether interaction of the tariff and bindings biases the uncertainty coefficient, column (4) of Table 4 keeps only observations with applied MFN tariffs of zero. This regression measures a pure delay effect since neither preferential market access or applied barriers in general would hold back entry. The uncertainty measure continues to have a negative effect.

4.5.2 Policy Evaluation

The exporter-year and exporter-product effects absorb a large share of total variation in the pattern of traded goods, nearly 0.8 in most of the regressions. This is not surprising given the well-known, traditional roles of technology driven comparative advantage, distance, exchange rates, transport costs and endowment differences in predicting the pattern of trade. Nevertheless, most of these factors cannot be directly influenced by trade policy, even over the long run. I focus on comparing the relative impacts of alternative trade policy instruments. I will show that in some scenarios the aggregate impact of trade policy uncertainty is substantial.

I use the econometric model to compare the scope for product growth given the margins of policy adjustment available to Australia during the sample period. I consider a policy experiment in which Australia jointly reduces all applied tariffs and bindings to zero in 2001. This is an informative benchmark year because it corresponds to the start date of the Doha Round of multilateral trade negotiations where bindings have been a contentious issue.

To quantify the impact of this policy change, I compute the total number of traded products as the sum of the changes in the predicted probability a product is traded over the entire sample in 2001, conditional on the fixed effects.[24] I use the actual levels of applied tariffs and bindings in the data. If a product already

[24]In the LPM this is equivalent to evaluating the estimation equation at the mean of the tariff and binding reductions and

has a tariff and binding of zero, then no new products will be predicted by the policy experiment.

The total change in the probability a product is traded can be decomposed into an applied tariff effect, binding lock-in effect, and a caution effect. First, when tariffs are decreased to zero, i.e. $\tau = 1$, the uncertainty measure will increase to $U^B(\tau = 1) = 1 - (B^{-\sigma}/1)$ as the gap between tariffs and and bindings widens. This will attenuate the impact of the tariff reduction. Second, initially suppose bindings are reduced to the level of applied tariffs so the $U^B(B_v = \tau_{tv}) = 1 - (\tau_{tv}/\tau_{tv}) = 0$. This will eliminate the incentive to delay entry, as tariffs are "locked-in" at present levels and can never be worse. Third, if bindings are further reduced to zero the attenuation of the initial tariff reduction is eliminated. Then the total effect in terms of the estimated coefficients can be computed for each exporter product and decomposed as follows

Decomposition of Total Effect for Policy Experiment

(A) Tariff Effect (lower tariffs to zero without binding):	$\mathrm{d}\ln\tau = \ln\tau_{tv,MFN} - 0$
$-b_\tau^* \mathrm{d}\ln\tau_{tv,MFN} + b_B^*[U_{\tau=1}^B - U_{tv}^B] - b_{pref}^* \ln(\tau_{tv,MFN}/\tau_{tjv})$	
(B) Binding Lock-in Effect (set bindings to level of tariff):	$\mathrm{d}\ln B = \ln B_{tv} - \ln\tau_{tv,MFN}$
$-b_B^* U_{tv}^B$	
(C) Remove Caution Effect (further binding tariffs at zero):	$\mathrm{d}\ln B = \ln\tau_{tv,MFN}$
$-b_B^*[U_{\tau=1}^B - U_{tv}^B]$	

I add up these changes over all exporter-product pairs in 2001 and report the results in Table 5. The model predicts a 4.1% increase in traded products if Australia were to set all its remaining positive MFN tariffs to zero on a unilateral basis. Alternatively, Australia could reduce all bindings to current applied tariffs, eliminating the risk of future "bad news." This action would increase the number of traded products by 8.8%. The remarkable aspect of this effect is that not a single *applied* policy measure would need to change. Merely the commitment to lock-in previous liberalization would generate a an increase in traded products. Lastly, bindings could be further reduced to zero in conjunction with the reduction of MFN tariffs to zero. This would generate an additional 4.8% increase from securing those tariff reductions so that the caution effect does not attenuate the response of entry.

These magnitudes of these predictions are conditional on the full set of fixed effects and would ultimately be sensitive to the underlying economic conditions. As such, it is informative to look at the share of growth for each policy change relative to the total growth achieved by setting bindings and tariffs to zero simultaneously. More than half of the potential new product growth is accounted for by removing uncertainty. Reducing bindings to applied tariffs accounts for half of the growth and removing caution accounts for an additional quarter of total growth. In contrast, the share of growth from unsecured, unilateral MFN tariff reductions

multiply the the sample size. A similar quantification exercise is used by Debaere and Mostashari (2010) in a different context.

is less than 25%.

Next, I consider a counterfactual scenario to the Uruguay Round. How many fewer products would have been traded if the Uruguay Round (UR) bindings had never been put in place, but Australia did continue to lower its applied MFN tariffs? I construct a counterfactual uncertainty measure holding binding commitments fixed at their Tokyo Round levels, but using the applied MFN tariff in 2001. Specifically, for each product I compute

$$U^{CF}_{t=2001,v} = 1 - \left(\frac{B_{1993,v}}{\tau_{2001,v,MFN}}\right)^{-\sigma} \tag{22}$$

for the subset of products that were bound in the Tokyo Round.

I then compute the increase in products in 2001 relative to the level predicted in 1993, at the mean tariff reduction and counterfactual uncertainty measure in (22). The counterfactual increase in the number of traded products reported in Table 6 is only 1% when bindings are held fixed over the time period. This anemic product growth is an economic and statistically significant 7% below the full effect of the UR binding and tariff reductions. To put this in terms of trade values, the second column of Table 6 uses the 2001 trade value weighted average of the predicted effect. The trade weighted difference in product growth is an even larger 11%.

While some predicted effects in these quantification exercises appear to be quite large, it is also possible these product measures actually understate the true level of entry by firms. There is undoubtedly within product firm entry. If a product is already traded or becomes traded due to the policy change, entry by more than one firm is counted only once. But whether the estimates over- or understate the true impact is less relevant when evaluating the relative efficacy of reducing unilateral applied tariffs versus reducing uncertainty. As long as the predictions are not systematically skewed toward applied protection or uncertainty, the relative contribution of uncertainty reductions are at least as effective as tariff reductions.

4.6 Robustness and Extensions

4.6.1 Robustness of LPM specification

One issue with the LPM is that it does not constrain the dependent variable to the unit interval, and therefore may suffer from heteroskedastic and non-normal errors. I cluster standard errors at the tariff line by year level in all regression to handle possible within tariff line correlation given the binding doesn't vary across countries. This is conservative, as alternatively clustering on the exporter-product panel identifier has lower standard errors in general on the regressors.

Another potential issue is that the estimated marginal effects are linear in continuous regressors, when the true distribution function of the underlying latent variable is non-linear. However, Angrist (2001) and Angrist and Pischke (2009) argue that the LPM typically does a very good job approximating the conditional expectation of the marginal effect on the probability of an event, and this is exact when the regressors are a fully saturated set of dummy variables. I construct two binary indicators, high tariffs and high binding gaps, equal to unity when a regressor is in the top third of its sample distribution.[25] The results in Table 7 include a full set of exporter-year and exporter-product fixed effects. The probability a product is traded is decreasing and significant for high tariff and high binding gap products. The estimates are robust to controlling for a positive preference margin indicator in column (2). In column (3), I take a subsample of products with zero MFN tariffs, in which case controlling for the preference margin is irrelevant, and continue to find a negative effect for high binding gap products.

Next, I run conditional logit regressions in Table 8 that confirm the sign and significance patterns of the LPM. The conditional logit partials out the exporter-product group effect to obtain consistent estimates of parameters. For computational purposes I control for exporter specific time trends rather than exporter-year effects. The practical cost of this estimation method is that coefficients for exporter-product groups that are always traded or never traded cannot be identified since there is no within group variation in the dependent variable. As the sample size shows, this drops a large number of exporter-product-year observations. If we want to make policy inferences from regressions about all of Australia's trade partners, the estimated marginal effects of the LPM are more representative.[26] Nevertheless, the signs and significance pattern from conditional logit regressions in columns (1) and (2) are consistent with the baseline LPM results. Columns (3) and (4) show the LPM results on the conditional logit subsample for comparison. Similar LPM results prevail using exporter-year effects rather than exporter time trends in columns (5) and (6).

4.6.2 Total Trade

As a final robustness check, I test whether trade policy uncertainty is also present in aggregate trade at the exporter-product level. The theoretical model makes predictions about the entry of firms into exporting assuming only that firm unit costs are bounded below by c_j^L in order generate zeros in exporter-product bilateral trade. To derive a log linear gravity equation, I must now assume unit costs have a Pareto distribution with unit costs bounded on $[c_j^L, c_j^H]$ such that $G(c) = \frac{c^k - (c_j^L)^k}{(c_j^H)^k - (c_j^L)^k}$ for the mass of firms in each exporter N_j. When the cutoff c_{tjv}^U is above the minimum unit cost for the exporter then total exports R_{tjv}

[25] The ranking of the binding to tariff ratio is the same whether I use the binding gap measure from the reduced form regression or the uncertainty measure in the structural specification.

[26] See Beck(2011) for further discussion of this shortcoming of the conditional logit relative to the LPM. In addition, only marginal effects that assume the unidentified exporter-product group effect is zero are feasible, and so they cannot be compared directly to OLS marginal effects.

are observed and equal to

$$R_{tjv} = \sigma A_{tv} \tau_{tv}^{-\sigma} N_j \int_{c_j^L}^{c_{tjv}^U} c^{1-\sigma} dG(c) \text{ if } c_{tjv}^U > c_j^L.$$

Evaluating the integral and taking logs yields

$$\begin{aligned}
\ln R_{tjv} &= -\sigma \ln \tau_{tv} + (k - \sigma + 1) \ln c_{tjv}^U + \ln \frac{k}{k - \sigma + 1} + \ln \left(\frac{1 - (c_j^L / c_{tjv}^U)^k}{(c_j^H)^k - (c_j^L)^k} \right) + \ln N_j + \ln \sigma + \ln A_{tv} \\
&= -\frac{-k\sigma}{\sigma - 1} \ln \tau_{tv} + \frac{(k - \sigma + 1)}{\sigma - 1} \ln \left[\frac{1 - \beta + \beta\gamma\Delta}{1 - \beta + \beta\gamma} \right] \\
&\quad + \ln \frac{k\sigma}{k - \sigma + 1} + \ln(1 - (c_j^L / c_{tjv}^U)^k) - \ln[(c_j^H)^k - (c_j^L)^k] + \ln N_j + \ln A_{tv} \qquad (23)
\end{aligned}$$

Using the first order approximation to the term in brackets from (17) and again controlling for the preference margin, I obtain an estimating equation

$$\ln R_{tjv} = b_B^R U_{tv}^B + b_\tau^R \ln \tau_{tv} + b_{pref}^R (\ln \tau_{tv,MFN} - \ln \tau_{tjv}) + d_{jt} + d_{jv} + \varepsilon_{tjv} \qquad (24)$$

The set of fixed effects are the same as the entry estimation equation (19), but the exporter-year effects d_{jt} now control for additional terms in $\ln N_j - \ln[(c_j^H)^k - (c_j^L)^k]$. These capture the mass of domestic firms in the exporting country, N_j, and unit cost heterogeneity in the pareto bounds. I have omitted the term $\ln[1 - (c_j^L / c_{tjv}^U)^k]$, which captures heterogeneity of entrants through the ratio of the pareto lower bound to the cutoff in each exporter-product. To understand how this term effects the estimated coefficients on uncertainty and tariffs, I differentiate it with respect to the cutoff

$$\frac{\partial \ln[1 - (c_j^L / c_{tjv}^U)^k]}{\partial \ln c_{tjv}^U} = (k - \sigma + 1) \frac{1}{(c_{tjv}^U / c_j^L)^{k - \sigma + 1} - 1} = \vartheta > 0.$$

Using ϑ to denote this derivative, the structural coefficients on any trade barriers are reinforced. The coefficients on binding uncertainty and tariffs are rescaled into total exports by the pareto distribution and the adjustment by ϑ. This means than any estimated coefficients are averaged over the values of ϑ, but have the same signs as the entry regressions: $b_B^R = -\gamma p_B \left[\frac{(k-\sigma+1)}{\sigma-1} + \vartheta \right] \frac{\beta}{1-\beta} < 0$ and $b_\tau^R = \frac{\sigma}{\sigma-1}(k + \vartheta) < 0$.[27]

There is a large prevalance of zero trade flows in the data when no firm has unit costs above the exporter-product cutoff. When estimating by OLS in log levels, all these zero trade flow observations are dropped. To account for this, I also employ the procedure of Santos Silva and Tenreyro (2006). They show that a poisson pseudo-maximum likelihood (PPML) regression can accommodate zeroes in the dependent variable

[27] In Helpman et al. (2008), a control function is used to handle this term in order to estimated the elasticity of *firm* sales to trade barriers.

and consistently estimate coefficients of a gravity equation. For comparison, I estimate equation (24) by both OLS, dropping all zero flows, and PPML in Table 9. Column (1) shows OLS results without controlling for the preference margin. The coefficients on tariffs and the uncertainty measure are negative and significant. There is little change when controlling for the preference margin in column (2). Columns (3) and (4) show the same regressions estimated via PPML. These regressions partial out the exporter-product effect and employ exporter trends instead of exporter-year fixed effects for computational stability. The sample size goes up substantially, as do the magnitudes of the coefficients on uncertainty and the applied tariff.[28]

As noted in section 4, the fixed effects probit regression is not consistent generally and computationally unfeasible in this application. This means the selection correction and control function procedure in Helpman et al. (2008) is not available. Moreover, I do not have a valid exclusion restriction as the uncertainty measure is significant in the gravity regression.[29] There are at least two explanations for this. First, multiple firms may enter within an exporter-product group at once or over time. This will increase total trade, but whether this is increase in trade volumes is from extensive or intensive adjustment at the firm-level is unobservable. Second, incumbent firms may make subsequent irreversible investment decisions that are also influenced by trade policy uncertainty (see for example Handley and Limão, 2013).

5 Conclusion

Trade policy is inherently uncertain. Multilateral policy commitments at the WTO are meant to provide a more secure and stable trade policy regime for prospective exporters. I capture these elements in a tractable model of export market entry patterns under uncertainty and test its predictions empirically. Evidence from Australia suggests that prospective exporters place weight on the possibility of trade policy reversals to bindings. This leads to delay of the entry decision and less responsiveness on the entry margin to trade policy changes. I find that multilateral policy commitments at the WTO help to reduce this uncertainty and increase product entry. Within the space of trade policy tools available, policy commitments could generate as more product entry than unilateral tariff reductions. These results are important for both quantifying the value and modeling the impact of tariff binding commitments at the WTO. The evidence of greater product entry in tariff lines with lower bindings, a key policy instrument for guaranteeing predictable market access, indicates that these commitments are valuable to exporters.

These findings point to an important and broad role for the WTO and international institutions in

[28]The fixed effect PPML estimator drops any exporter-product group with all zero outcomes, so I do not recover the full set of observations in the LPM entry panel.

[29]Helpman et al. (2008) decompose their export equation into intensive and extensive components. This allows for consistent estimation in a cross-section of the effect of distance and other covariates on parameters of the *firm-level* sales equation. Here, the set of panel and exporter-year fixed effects account for some of the heterogeneity in productivity with the rest of the impact going into the rescaled coefficients on tariffs and uncertainty.

the monitoring and enforcement of multilateral commitments. For bindings to be effective, exporters must believe that penalties for violating the commitments would be costly enough that they would either never occur or be swiftly reversed. The evidence that lower bindings increase product entry for Australia suggests they are credible, but this might vary by country. To that extent, the WTO could play an important role in monitoring and enforcement of multilateral agreements on services, investment and intellectual property.

Extending and verifying these results with a broader group of countries and applications outside of international trade is important. Fortunately, the methodology developed here, which uses product data and model-based policy uncertainty measures, can be applied more broadly within international trade applications and to other forms of policy uncertainty. An important extension is the impact of trade policy uncertainty on foreign direct investment where sunk costs of opening a production facility may be even higher. Trade policy uncertainty takes many other forms in the world trade system. Modeling and testing the risk of non-renewal in preferential tariff programs, temporary trade bans, economic sanctions and the risk of anti-dumping measures are all subjects for future work.

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A Appendix

A.1 CDF of bound tariff distribution

The observed tariff in the bound regime τ_B is censored at the binding rate of B.

$$\tau_B = \begin{cases} \tau & \text{if } \tau \le B \\ B & \text{if } \tau > B. \end{cases}$$

The CDF of τ_B is $H_{\tau_B}(\tau_B) = pH_1(\tau) + (1-p)H_2(\tau)$. Where $p = H(B)$ and

$$H_1(\tau) = \begin{cases} \frac{H(\tau)}{H(B)} & \text{if } \tau \le B \\ 1 & \text{if } \tau > B, \end{cases} \quad \text{and } H_2(\tau) = \begin{cases} 0 & \text{if } \tau \le B \\ 1 & \text{if } \tau > B. \end{cases}$$

A.2 Value functions and stochastic cutoff condition

Four equations define the initial problem at time t of a firm with unit labor requirement c. For clarity of exposition, I drop the country of origin subscripts. The value of exporting is

$$V^1(\tau_t) = \pi(\tau_t) + \beta[\underbrace{(1-\gamma)V^1(\tau_t)}_{\text{No Shock}} + \underbrace{\gamma EV^1(\tau')}_{\text{Shock Arrives}}]. \tag{25}$$

Following a policy shock, the *ex-ante* expected value of exporting following a policy change to a new tariff τ' is

$$EV^1(\tau') = E\pi(\tau') + \beta[(1-\gamma)EV^1(\tau') + \gamma EV^1(\tau')] \tag{26}$$

In time period t, the unconditional expected value of being an exporter next period given that a policy shock arrives is $EV^1(\tau')$ in (26). This expectation is time invariant because I assume that the distribution of future tariffs $H(\tau')$ is time invariant. Equation (26) can be solved explicitly for $EV^1(\tau')$ to obtain

$$EV^1(\tau') = \frac{E\pi(\tau')}{1-\beta}.$$

The resulting time invariance of $EV^1(\tau')$ does not mean that the value of exporting is time invariant. $V^1(\tau_t)$ is a function of the current tariff and firms can re-compute it on an *ex-post* basis following every tariff policy change.

The value of waiting is

$$V^0(\tau_t) = 0 + \beta[\underbrace{(1-\gamma)V^0(\tau_t)}_{\text{No Shock}} + \underbrace{\gamma(1-H(\tau_1))V^0(\tau_t)}_{\text{Shock Above Trigger}} + \underbrace{\gamma H(\tau_1)(EV^1(\tau_1 \mid \tau \le \tau_1) - K_e)}_{\text{Shock Below Trigger}}] \tag{27}$$

Conditional on waiting until the tariff falls below the entry trigger, the expected value of exporting is

$$EV^1(\tau_1 \mid \tau \le \tau_1) = E\pi(\tau' \mid \tau' < \tau_1) + \beta[(1-\gamma)EV^1(\tau_1 \mid \tau_t \le \tau_1) + \gamma EV^1(\tau')] \tag{28}$$

This equation is structurally the same as (25), but it is evaluated *ex-ante* to obtain the expected value of exporting to a firm that delays entry until a more favorable policy shock arrives. If a firm waits to enter in the current period, it must be the case that current tariff exceeds the entry trigger, $\tau_t > \tau_1$.

The set of four equations (25),(26),(27), and (28) is a linear system in the four quantities $V^1(\tau_t)$,$EV^1(\tau')$,

$V^0(\tau_t)$, and $E[V^1(\tau_1 \mid \tau \leq \tau_1)]$ and can be solved explicitly. The solution to the set of equations is

$$V^1(\tau_t) = \frac{\pi(\tau_t)[1-\beta(1-\gamma E[\pi(\tau')])]}{[1-\beta(1-\gamma)](1-\beta)} \tag{29}$$

$$EV^1(\tau') = \frac{E[\pi(\tau')]}{1-\beta} \tag{30}$$

$$V^0(\tau_t) = \beta\gamma H(\tau_1)\frac{(1-\beta)E[\pi(\tau') \mid \tau' < \tau_1] - \beta\gamma E[\pi(\tau')] - (1-\beta)[1-\beta(1-\gamma)]K_e}{[1-\beta(1-\gamma)][1-\beta(1-\gamma H(\tau_1))]} \tag{31}$$

$$EV^1(\tau_1 \mid \tau < \tau_1) = \frac{\beta\gamma E[\pi(\tau')] - E[\pi(\tau') \mid \tau' < \tau_1](1-\beta)}{(1-\beta+\beta\gamma)(1-\beta)} \tag{32}$$

The entry margin corresponds to the firm with unit labor requirement c^U that is just indifferent to entry or waiting at time t. For this marginal firm, the current tariff equals the entry trigger such that $\tau_t = \tau_1$. Equations (29) and (31) for the value of exporting today versus waiting can be substituted into indifference condition (6) to obtain the cutoff condition in the main text.

A.3 Profit Loss Term $\Delta(\tau_t)$

A.3.1 $\Delta(\tau_t) \leq 1$

I denote the maximum tariff by τ_{max}.

$$\begin{aligned}
\Delta(\tau_t) &= \left[E(\tau^{-\sigma}) + H(\tau_t)[\tau_t^{-\sigma} - E(\tau^{-\sigma} \mid \tau \leq \tau_t)]\right]/\tau_t^{-\sigma} \\
&= \left[\int_1^{\tau_{max}} \tau^{-\sigma} dH(\tau) + H(\tau_t)\tau_t^{-\sigma} - \int_1^{\tau_t} \tau^{-\sigma} dH(\tau)\right]/\tau_t^{-\sigma} \\
&= \left[\int_{\tau_t}^{\tau_{max}} \tau^{-\sigma} dH(\tau) + H(\tau_t)\tau_t^{-\sigma}\right]/\tau_t^{-\sigma} \\
&= \left[(1-H(\tau_t))E(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma}\right]/\tau_t^{-\sigma}
\end{aligned}$$

Then to show that $\Delta(\tau_t) \leq 1$, I take the difference D of the numerator and denominator in the final line above

$$\begin{aligned}
D &= \left[(1-H(\tau_t))E(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma}\right] - \tau_t^{-\sigma} \\
&= (1-H(\tau_t))[E(\tau^{-\sigma} \mid \tau \geq \tau_t) - \tau_t^{-\sigma}] \\
&\leq 0
\end{aligned}$$

The inequality follows because $\tau_t^{-\sigma}$ is always greater than $E(\tau^{-\sigma} \mid \tau > \tau_t)$. 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.

A.3.2 Derivation of $\Delta(\tau_t, B)$ when tariffs are bound.

$$\begin{aligned}
\Delta(\tau_t, B) &= \left[E(\tau^{-\sigma}) + H(\tau_t)[\tau_t^{-\sigma} - E(\tau^{-\sigma} \mid \tau \leq \tau_t)]\right]/\tau_t^{-\sigma} \\
&= \left[(1-H(B))B^{-\sigma} + \int_1^{B} \tau^{-\sigma} dH(\tau) + H(\tau_t)\tau_t^{-\sigma} - \int_1^{\tau_t} \tau^{-\sigma} dH(\tau)\right]/\tau_t^{-\sigma} \\
&= \left[(1-H(B))B^{-\sigma} + \int_{\tau_t}^{B} \tau^{-\sigma} dH(\tau) + H(\tau_t)\tau_t^{-\sigma}\right]/\tau_t^{-\sigma} \\
&= \frac{(1-H(B))B^{-\sigma} + [H(B) - H(\tau_t)]E(\tau^{-\sigma} \mid \tau_t < \tau < B) + H(\tau_t)\tau_t^{-\sigma}}{\tau_t^{-\sigma}}
\end{aligned} \tag{33}$$

A.4 Proofs of Propositions 1 and 2

PROPOSITION 1 ***[Delay]*** *Higher bindings reduce the entry cutoff, c^U, by delaying investment in market entry, and making entry more sensitive to the policy shock arrival rate. In elasticity terms:*

(a) The cutoff is decreasing in the binding

$$\varepsilon(B) = \frac{d\ln c_t^U}{d\ln B} = -\frac{\sigma}{\sigma-1}\left(\frac{\beta\gamma}{(1-\beta+\beta\gamma\Delta(\tau_t,B))}\left(\frac{(1-H(B))B^{-\sigma}}{\tau_t^{-\sigma}}\right)\right) < 0$$

(b) The cutoff is decreasing arrival rates

$$\varepsilon(\gamma) = \frac{d\ln c_t^U}{d\ln\gamma} = \frac{\beta\gamma}{\sigma-1}\left[\frac{1-\beta}{(1-\beta+\beta\gamma)[1-\beta+\beta\gamma\Delta(\tau_t,B)]}\right][\Delta(\tau_t,B)-1] \leq 0$$

(c) The cutoff is more responsive to arrival rates when bindings increase $\frac{\partial\varepsilon(\gamma)}{\partial B} = \frac{d^2\ln c_t^U}{d\ln\gamma\cdot dB} \leq 0$

PROOF:

(a) I use the binding censored version of the profit loss term $\Delta(\tau_t, B)$. Log differentiating the cutoff, I obtain

$$\begin{aligned}\frac{d\ln c_t^U}{d\ln B} &= \frac{1}{\sigma-1}\left(\frac{\beta\gamma}{(1-\beta+\beta\gamma\Delta(\tau_t,B))}\frac{d\Delta(\tau_t,B)}{d\ln B}\right)\\ &= -\frac{\sigma}{\sigma-1}\left(\frac{\beta\gamma}{(1-\beta+\beta\gamma\Delta)}\left[\frac{(1-H(B))B^{-\sigma}}{\tau_t^{-\sigma}}\right)\right] \ < \ 0.\end{aligned}$$

The term is brackets is positive and the cutoff is decreasing in the binding. ■

(b) Log differentiating the cutoff under uncertainty with respect to γ, I obtain

$$\begin{aligned}\frac{d\ln c_t^U}{d\ln\gamma} &= \frac{\gamma}{\sigma-1}\left(\frac{d}{d\gamma}\ln(1-\beta\,(1-\gamma\Delta(\tau_t,B))) - \frac{d}{d\gamma}\ln\,(1-\beta(1-\gamma))\right)\\ &= \frac{\beta\gamma}{\sigma-1}\left[\frac{1-\beta}{(1-\beta+\beta\gamma)(1-\beta+\beta\gamma\Delta(\tau_t,B))}\right][\Delta(\tau_t,B)-1] \qquad (34)\end{aligned}$$

We thus have

$$\text{sgn}\left(\frac{d\ln c_t^U}{d\gamma}\right) = \text{sgn}\left(\frac{\Delta(\tau_t,B)-1}{(1-\beta+\beta\gamma\Delta(\tau_t,B))}\right) < 0$$

which is negative since $\Delta(\tau_t, B) - 1 < 0$ whenever $\tau_t < \tau_{max}$. ■

(c) The profit loss term is decreasing in B since $\frac{\partial\Delta(\tau_t,B)}{\partial B} = -\sigma\frac{(1-H(B))B^{-\sigma-1}}{\tau_t^{-\sigma}} < 0$. Then terms in $\Delta(\tau_t, B)$ in the denominator of the first term in brackets from (34) are decreasing. Thus the term in brackets is increasing. The second term $[\Delta(\tau_t, B) - 1]$ is decreasing. Therefore, $\frac{\partial\varepsilon(\gamma)}{\partial B} \leq 0$. ■

PROPOSITION 2 ***[Caution]*** *The entry cutoff c^U is less elastic with respect to a given tariff change under uncertainty ($\gamma > 0$).*

$$\begin{aligned}\varepsilon^U(\tau_t) &= \frac{d\ln c_t^D}{d\ln\tau_t} + \frac{d\ln\Theta_t}{d\ln\tau_t}\\ &= -\frac{\sigma}{\sigma-1}\left[1-\left(\beta\gamma\frac{(1-H(B))B^{-\sigma}+(H(B)-H(\tau_t))E[\tau^{-\sigma}\mid\tau\in(\tau_t,B)]}{[1-\beta+\beta\gamma\Delta(\tau_t,B)]\tau_t^{-\sigma}}\right)\right].\end{aligned}$$

PROOF:

As described in the main text, the proof consists of two parts. First, I show that the expected profit loss of a bad shock is decreasing in the current tariff τ_t. Second, I show the stochastic elasticity is proportionally less than the deterministic elasticity.

(1) $\frac{\partial \Delta(\tau_t,B)}{\partial \tau_t} \geq 0$ implies the proportion of profits lost in a tariff reversal, $\Delta(\tau_t, B) - 1$, is reduced as tariffs increase. In semi-elasticity terms, this is

$$\begin{aligned}
\frac{\partial \Delta(\tau_t, B)}{\partial \tau_t}\tau_t &= \tau_t[-\tau_t^{-\sigma}h(\tau_t) + h(\tau_t)\tau_t^{-\sigma} - \sigma H(\tau_t)\tau_t^{-\sigma-1}]/\tau_t^{-\sigma} \\
&\quad + \tau_t[(1 - H(\tau_t))E(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma}](\sigma\tau_t^{\sigma-1}) \\
&= [-\sigma H(\tau_t)]/\tau^{-\sigma} + \sigma[(1 - H(\tau_t))E(\tau^{-\sigma} \mid \tau \geq \tau_t) + H(\tau_t)\tau_t^{-\sigma}]/\tau_t^{-\sigma} \\
&= \sigma[(1 - H(\tau_t))E(\tau^{-\sigma} \mid \tau \geq \tau_t)]/\tau_t^{-\sigma} \\
&= \sigma[(1 - H(B))B^{-\sigma} + (H(B) - H(\tau_t))E(\tau^{-\sigma} \mid \tau_t \leq \tau < B)]/\tau_t^{-\sigma} \\
&= \frac{\partial \Delta(\tau_t, B)}{\partial \ln \tau_t} \geq 0
\end{aligned}$$

(2) Using the expression for c_j^U from in equation (7), I log differentiate and derive the elasticity

$$\begin{aligned}
\varepsilon^U(\tau_t) &= \frac{d \ln c_t^D}{d \ln \tau_t} + \frac{d \ln \Theta_t}{d \ln \tau_t} \\
&= -\frac{\sigma}{\sigma - 1} + \frac{1}{\sigma - 1}\left(\frac{\beta\gamma}{(1 - \beta + \beta\gamma\Delta(\tau_t, B))}\frac{d\Delta(\tau_t, B)}{d \ln \tau_t}\right) \\
&= -\frac{\sigma}{\sigma - 1}\left[1 - \left(\beta\gamma\frac{(1 - H(B))B^{-\sigma} + (H(B) - H(\tau_t))E[\tau^{-\sigma} \mid \tau \in (\tau_t, B)]}{[1 - \beta + \beta\gamma\Delta(\tau_t, B)]\tau_t^{-\sigma}}\right)\right] \\
&= -\frac{\sigma}{\sigma - 1} \times \phi(\tau_t) \\
&= \varepsilon^D(\tau) \times \phi(\tau_t)
\end{aligned}$$

The term in brackets, represented by $\phi(\tau_t)$, is less than or equal to one. Therefore, in absolute values $\mid \varepsilon^U(\tau_t) \mid < \mid \varepsilon^D(\tau_t) \mid$. ■

A.5 Data Sources and Descriptions

I use trade flow and product data for all imported exporter-product pairs from 1991 to 2001. These data are at 10-digit level of disaggregation known as the Harmonized Tariff Items Statistical Codes (HTISC) by Australian Customs. The data were obtained on an annual basis from Trade Data International, an authorized re-seller of trade data from the Australian Bureau of Statistics.[30] After aggregating to the 8 digit level and concording over time with the tariff data, there are 4,673 products that could be exported from any single country to Australia. There are 630,702 exporter-year fixed effects in the final sample.

Applied MFN tariffs and preferences were obtained from the the UN TRAINS system via download from the the World Integrated Trade Solution (WITS). Uruguay Round bindings at the 8 digit level were extracted from the WTO's Tariff Analysis On-line system, a comprehensive database tariff concessions. The Consolidated Tariff Schedules contain a record of Australia's certified binding concessions.

The Tokyo Round bindings were negotiated at the conclusion of the the round in 1979 and fully implemented by 1987. I obtained a CD-ROM from the USITC with the 8 digit binding commitments from Version 1.0 of the Integrated Database, optimized for Windows 95. This CD could only be run on a 32-bit machine running Windows-XP software. The software program allows only one tariff line at time to be viewed. Each line was captured and exported to a text file that was then parsed and read into a cleaned up dataset of Tokyo Round bindings for analysis.

[30]The HTISC is equivalent to the Harmonized System in the first 6 digits, known as HS6 level. Following the HS6, the next 2 digits capture "tariff items" and are assigned for further disaggregation of tariff duties. The final 2 digits are "statistical codes" assigned to provide additional disaggregation for statistical purposes.

GATT/WTO membership data were obtained from the CEPII "Gravity" dataset available online. The import demand elasticities estimated by Kee et al. (2009) were obtained from the World Bank website.

A.6 Balanced Panel Results

The results in the main text use an unbalanced panel either because products are unbound or because the membership of the GATT/WTO changes over time. Table A1 reports results on the the Tokyo Round bound subsample of exporter-products in column (1). In column (2) the sample is restricted further to the set of countries that were GATT/WTO members from 1991-2006. The column (2) results are a true balanced panel. Column (3) is a balanced panel restricted to the years following the Uruguay Round, 1996-2001 when all exporters are WTO members for the entire period. The main conclusions using an unbalanced panel are unchanged.

Table A1: Pre- and Post-Uruguary Round sample split – balanced and unbalanced

Dependent Variable: Product Traded(binary)				
	(1)	(2)	(3)	(4)
Sample:	TR Bound ('91-'01)		Post UR ('96-'01)	
	Unbalanced	Balanced	Unbalanced	Balanced
Binding Uncertainty	-0.0291***	-0.0304***	-0.0453***	-0.0460***
	[0.00370]	[0.00388]	[0.0139]	[0.0141]
Applied Tariff (ln)	-0.256***	-0.265***	-0.160***	-0.162***
	[0.0288]	[0.0299]	[0.0408]	[0.0413]
Preference Margin	0.294***	0.306***	0.132***	0.135***
	[0.0512]	[0.0541]	[0.0481]	[0.0490]
Observations	917,602	739,288	3,575,580	3,444,876
R-squared	0.79	0.788	0.804	0.802

Notes: All columns include exporter-product and exporter-year fixed effects. Robust standard errors in brackets clustered by product-year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. $\sigma = 4$ for construction of binding uncertainty measure.

Table 1: Summary Statistics for Baseline Sample

	Mean	**St. Dev.**	**Min.**	**Max.**	**Obs.**
Product Traded(binary)	0.068	0.252	0.000	1.000	3770862
Exports to Australia(ln)	12.045	1.890	9.210	21.856	256084
Exports (1000s)	132.83	6111.30	0	3100000	3770862
Applied Tariff(ln)	0.045	0.064	0.000	0.315	3770862
Binding Gap (ln)	0.049	0.055	0.000	0.438	3770862
Binding (ln)	0.094	0.098	0.000	0.438	3770862
Binding Uncertainty ($\sigma = 4$)	0.162	0.160	0.000	0.827	3770862
Preference Margin	0.004	0.012	0.000	0.223	3770862
Binding Unc (sectoral σ)	0.057	0.064	0.000	0.514	3765500
Binding Unc (HS2 σ)	0.095	0.128	0.000	0.972	3730000

Table 2: Probability a product is traded – Reduced Form Regressions

Dependent Variable: Product Traded(binary)	(1)	(2)	(3)
Binding Gap(ln)	-0.0804***	-0.0915***	
	[0.00819]	[0.00888]	
Applied Tariff (ln)	-0.135***	-0.164***	-0.0942***
	[0.0101]	[0.0126]	[0.0107]
Preference Margin		0.143***	0.0371
		[0.0289]	[0.0267]
Observations	3,770,862	3,770,862	3,770,862
R-squared	0.796	0.796	0.796

Notes: All columns include exporter-year and exporter-product fixed effects. Robust standard errors in brackets clustered by product-year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 3: Probability a product is traded – Baseline Regressions

Dependent Variable: Product Traded(binary)						
	(1)	(2)	(3)	(4)	(5)	(6)
Sample:	Baseline	Baseline	Baseline	Baseline	Preference Margin=0	Drop Pref. Tariff Lines
Binding Uncertainty	-0.0501***	-0.0373***	-0.0357***	-0.0319***	-0.0468***	-0.0559***
	[0.00290]	[0.00322]	[0.00322]	[0.00300]	[0.00464]	[0.00809]
Applied Tariff (ln)	-0.275***	-0.181***	-0.169***	-0.140***	-0.209***	-0.262***
	[0.0123]	[0.0129]	[0.0128]	[0.0103]	[0.0169]	[0.0285]
Preference Margin	0.211***	0.204***	0.143***			
	[0.0283]	[0.0289]	[0.0287]			
Fixed Effects:						
exporter-product	yes	yes	yes	yes	yes	yes
exporter-year	no	no	yes	yes	yes	yes
year	no	yes	N.A.	N.A.	N.A.	N.A.
Observations	3,770,862	3,770,862	3,770,862	3,770,862	3,209,898	791,403
R-squared	0.795	0.795	0.796	0.796	0.8	0.807

Notes: All columns include exporter-product fixed effects. Robust standard errors in brackets clustered by product-year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. $\sigma = 4$ for construction of binding uncertainty measure.

Table 4: Robustness to alternative parameterization of elasticity of substitition (σ) for uncertainty measure.

Dependent Variable: Product Traded(binary)				
	(1)	(2)	(3)	(4)
Elasticity parameter:	$\sigma = 4$	HS2 Level	Section Level	$\tau_{MFN} = 0$
Binding Uncertainty	-0.0357***	-0.0500***	-0.0723***	-0.0305***
	[0.00322]	[0.00486]	[0.00769]	[0.00576]
Applied Tariff (ln)	-0.169***	-0.168***	-0.166***	
	[0.0128]	[0.0136]	[0.0136]	
Preference Margin	0.143***	0.140***	0.146***	
	[0.0287]	[0.0293]	[0.0295]	
Elasticities of marginal probability a product is traded computed at mean of Binding Uncertainty Measure				
Applied Tariff	-0.049***	-0.019***	-0.074***	
	[0.011]	[0.012]	[0.011]	
Binding	-0.120***	-0.149***	-0.092***	-0.102***
	[0.011]	[0.014]	[0.010]	[0.031]
Cautionary Effect Attenuation (%)	70.8	88.7	55.3	N.A.
Observations	3,770,862	3,730,000	3,765,500	1,703,110
R-squared	0.796	0.796	0.796	0.777

Notes: All columns include exporter-year and exporter-product fixed effects. Robust standard errors in brackets clustered by product-year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Tariff and binding elasticity computed at mean of U^B_{tv} using equation (20), $e(\tau) = b_\tau - b_B\sigma \times (1 - U^B_{tjv})$, and (21), $e(B) = -b_B\sigma \times (1 - U^B_{tjv})$. See text for derivations. Mean of U^B_{tv} specific to each measure used in each column except for column (4), which uses overall sample mean from column (1) for comparison.

Table 5: Quantification of the role of policy uncertainty – predicted growth in traded products from policy changes in 2001.

Predicted new product growth rates and growth shares in 2001		
	Growth Rate	Share
Policy Experiments:		
A. Tariff Effect - reduce applied tariffs to zero	4.12	0.23
	[0.550]	
B. Binding Lock-in Effect - reduce binding to level of applied tariff	8.83	0.50
	[0.797]	
C. Caution Effect - jointly reduce bindings and applied tariffs to zero	4.81	0.27
(net of A and B)	[0.434]	
Total Effect	17.76	1.00
	[1.280]	

Notes: Estimates computed from coefficient estimates in column 3 of table 3 under the assumption that $\sigma = 4$ for uncertainty measure. Predicted number of products computed as the sum across all observations of the increase in the probability a product is traded by experiment and year. Totals do no add precisely due to rounding error. Growth rates computed relative to true number of traded products in the sample. All calculations are statistically different from zero with standard errors in brackets computed via the delta method. See text for exact formulas.

Table 6: Counterfactual quantification of product growth from 1993-2001 if bindings held fixed at 1993 levels.

	Growth Rate	
	Products	Trade Weighted Products
Predicted with full model	8.28	14.5
	[0.653]	[1.020]
Predicted without binding reductions	1.04	3.4
	[0.468]	[0.501]
Difference	7.24	11.1
	[0.653]	[1.000]

Notes: Estimates computed from coefficient in column 3 of table 3 under the assumption that $\sigma = 4$ for uncertainty measure. Only the subsample of tariff lines that were already bound by 1993 after the Tokyo Round are used. Predicted number of products computed as the sum across all observations of the increase in the probability a product is traded by experiment and year. Totals do no add precisely due to rounding error. Growth rates computed relative to true number of traded products in 1993. Trade weighted estimates use the value of trade in 2001. All calculations are statistically different from zero with standard errors in brackets computed via the delta method. See text for exact formulas.

Table 7: Robustness to Fully Saturated Linear Probability Model

Dependent Variable: Product Traded(binary)	(1)	(2)	(3)
High Uncertainty(binary)	-0.00212***	-0.00211***	-0.00453***
	[0.000570]	[0.000571]	[0.00173]
High Tariff (binary)	-0.0129***	-0.0139***	
	[0.00154]	[0.00211]	
Preference Margin(binary)		0.00144	
		[0.00129]	
Observations	3,770,862	3,770,862	1,703,110
R-squared	0.796	0.796	0.777

Notes: All columns include exporter-year and exporter-product fixed effects. Robust standard errors in brackets clustered by product-year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 8: Probability a product is traded – Conditional Logit Specifications and Sub-sample.

Dependent Variable: Product Traded(binary)	(1)	(2)	(3)	(4)	(5)	(6)
Specification:	c. logit	c. logit	LPM	LPM	LPM	LPM
Binding Uncertainty	-1.174***	-1.270***	-0.158***	-0.181***	-0.212***	-0.220***
	[0.112]	[0.118]	[0.0205]	[0.0216]	[0.0260]	[0.0266]
Applied Tariff (ln)	-4.112***	-4.632***	-0.647***	-0.787***	-0.779***	-0.848***
	[0.373]	[0.415]	[0.0804]	[0.0880]	[0.0883]	[0.0956]
Preference Margin		3.243***		0.802***		0.451**
		[1.120]		[0.187]		[0.216]
exporter time trends	yes	yes	yes	yes	no	no
exporter-year effects	no	no	no	no	yes	yes
Observations	274,479	274,479	274,479	274,479	274,479	274,479
R-squared			0.271	0.271	0.275	0.275

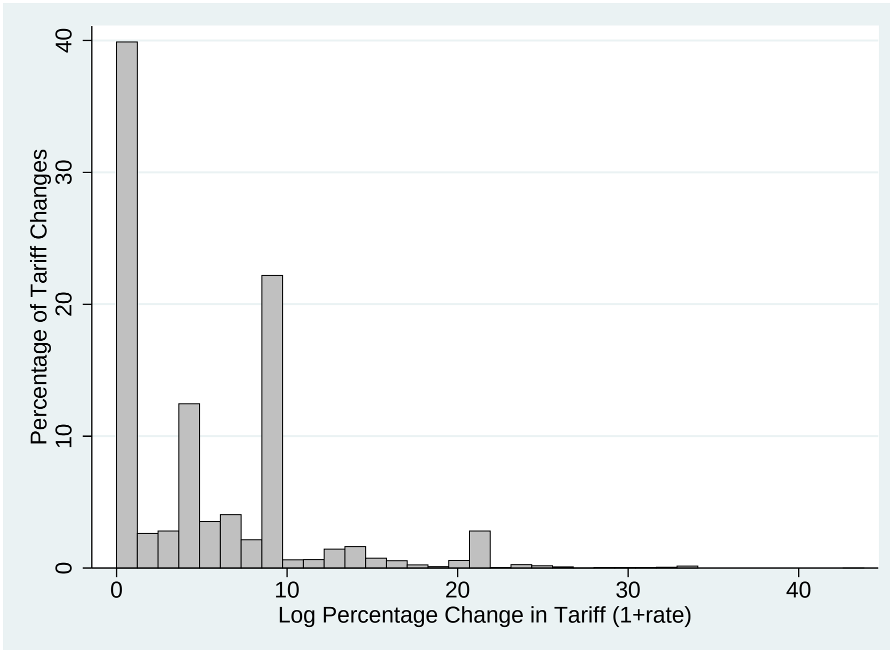
Notes: All columns include exporter-year and exporter-product fixed effects. Robust standard errors in brackets clustered by product-year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Coefficients for conditional logits in columns (1) and (2) are not transformed into marginal effects. Sample size reduced because conditional logit only identified for within group variation of dependent variable.

Table 9: Gravity regressions for total exports.

Dependent Variable:	Imports(ln)		Imports (levels)	
	(1)	(2)	(3)	(4)
	OLS	OLS	PPML	PPML
Binding Uncertainty	-0.460***	-0.451***	-0.355**	-0.342**
	[0.0756]	[0.0766]	[0.166]	[0.165]
Applied Tariff (ln)	-1.135***	-1.068***	-1.907***	-1.834**
	[0.243]	[0.251]	[0.676]	[0.721]
Preference Margin		-0.55		-0.824
		[0.586]		[1.241]
exporter time trends	no	no	yes	yes
exporter-year effects	yes	yes	no	no
Observations	256,084	256,084	417,592	417,592
R-squared	0.866	0.866	N.A.	N.A.

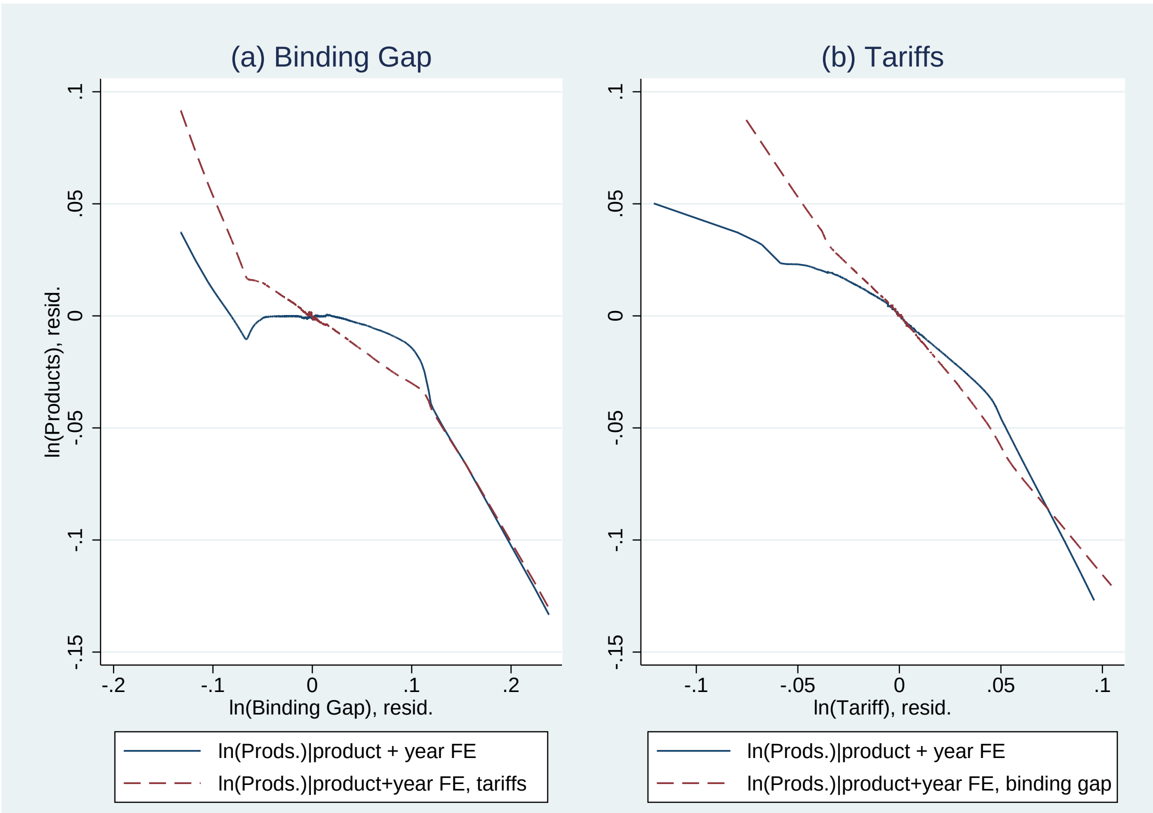
Notes: All columns include exporter-year and exporter-product fixed effects. Robust standard errors in brackets clustered by product-year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure 1: Distribution of tariff changes under a binding reversal in 2001



Notes: Change in log points from the MFN tariff to the bound tariff in 2001. $100 \times \ln(B_v/\tau_v)$ where $B, \tau = (1 + \text{ad-valorem rate})$.

Figure 2: Non-parametric relationship between number of traded products and: (a) Binding Gap, and (b) Tariffs.



Notes: Lowess curve of relationship between log product counts within each tariff line across all countries, conditional on product and year fixed effects and binding gap and tariffs, respectively.