Policy Uncertainty, Trade and Welfare:

Theory and Evidence for China and the U.S.*

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ABSTRACT: We examine the impact of policy uncertainty on trade and real income through firms’ export investments in general equilibrium. We quantify the trade policy causes and welfare effects of China’s export boom toward the U.S. following its 2001 WTO accession; reducing the threat of a trade war explains 22% of export growth to the U.S. Reduced policy uncertainty lowered U.S. prices, increasing consumers’ income by at least 0.8 percent, the welfare equivalent of an 8 percentage point tariff decrease. These findings provide evidence of large effects of policy uncertainty on economic activity and the importance of agreements in reducing it.

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1 Introduction

One of the most important economic developments of the last 20 years is China’s integration into the global trading system. The world’s share of imports from China between 1990-2010 rose from 2 to 11 percent and that increase was even larger for the U.S. (from 3 to 19 percent). This has translated into a more than tenfold increase in the share of US manufacturing expenditure on Chinese goods and there is evidence that this has contributed to declines in both U.S. prices (cf. Auer and Fischer, 2010) as well as manufacturing employment and local wages (cf. Autor et al., 2013). Figure 1 shows that most of this trade boom occurred after China’s accession to the World Trade Organization (December 2001), which has led some authors to argue that the accession may have reduced trade costs faced by Chinese exporters. This argument is somewhat puzzling given that U.S. applied trade barriers toward China remained largely unchanged at the time of accession.

We argue that China’s WTO accession significantly contributed to its export boom to the U.S. through a reduction in U.S. trade policy uncertainty. Specifically, China obtained permanent most favored nation (MFN) status with accession, which ended the annual U.S. threat to impose high tariffs. Had MFN status been revoked the U.S. would have reverted to Smoot-Hawley tariff levels and a trade war may have ensued. In 2000 for example, the average U.S. MFN tariff was 4% but if China had lost its MFN status it would have faced an average tariff of 31%. After WTO accession, the Chinese Foreign Trade Minister pointed out that by establishing “the permanent normal trade relationship with China, [the U.S.] eliminated the major long-standing obstacle to the improvement of Sino-U.S. (...) economic relations and trade.”

To examine this argument we build a model that allows us to interpret and measure the effects of trade policy uncertainty (TPU). We obtain structural estimates of key policy uncertainty parameters and use them to quantify the implications for aggregate prices and welfare of U.S. consumers. For example, we construct counterfactuals such as Chinese import penetration if U.S. policy uncertainty remained unchanged—shown in Figure 1. We also estimate a consumer welfare increase of 0.8% from reducing uncertainty that is equivalent to a 8 percentage point reduction in tariffs. Our approach and results have important implications beyond this specific event; they contribute to the growing literature on the impact of economic policy uncertainty, which we discuss below.

Our results formalize and quantify the impacts of TPU on investment and prices that policymakers and business leaders highlighted as important. For example, the U.S. decision to delink MFN from China’s human rights was described as having “removed a major issue of uncertainty” and that the impact of renewal on investment and re-exports “will remove the threat of potential losses that would have arisen as a result of revocation.” U.S. business leaders claimed that “…the imposition of conditions upon the renewal of MFN as virtually synonymous with outright revocation. Conditionality means uncertainty.” and lobbied Congress to make MFN permanent (Zeng, 2003). At the same time congressional research reports highlighted the

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1 Autor et al.(2013) make this point and also cite other motives for this export growth. China’s income has risen driven by internal reforms (many in the 1990s) with a subset targeted to exports (Hsieh and Klenow, 2009; Blonigen and Ma, 2010).
2 Although China never lost its temporary MFN status after it was granted in 1980, it came close: after the Tiananmen square protests there was pressure to revoke MFN status with Congress voting on such a bill every year in the 1990s and the House passing it three times.
4 “HK business leaders laud US decision” South China Morning Post, 5/28/94, Business section. The uncertainty recurred several times until the WTO accession. For example, in 1997 the Chinese Foreign Trade Minister urged the U.S. to abandon trade status reviews: “The question of MFN has long stymied the development of Sino-U.S. economic ties and trade (...)” [It] has created a feeling of instability among the business communities of the two countries and has not been conducive to bilateral trade development.” “Minister urges USA to abandon trade status reviews” Xinhua news agency, 10/5/97, FE/D3044/G.
5 Tyco Toys CEO “China MFN Status,” Hearing before the Committee on Finance, U.S. Senate, June 6, 1996, p. 97.
higher consumer prices that would result if MFN was ever revoked (Pregelj, 2001).

Our model captures the interaction between uncertainty and investment by modelling the latter as sunk costs and thus generating an option value of waiting. This basic theoretical mechanism is well understood (cf. Bernanke, 1983; Dixit, 1989), and there is some evidence that economic uncertainty, as proxied by stock market volatility, leads firms to delay investments (Bloom et al., 2007). In the international trade context, there is evidence of sunk costs to export market entry (cf. Roberts and Tybout, 1997), but most empirical research on uncertainty’s impact on export dynamics has focused on exchange rate uncertainty and finds small or negligible impacts (IMF, 2010). In a general equilibrium setting, Impullitti et al. (2013) find a sunk cost model with heterogeneous firms and uncertain efficiency fits observed aggregate trade dynamics well.

Much less is known about the implications of economic policy uncertainty. Early theoretical contributions to this issue (cf. Rodrik, 1991) recognized the difficulty in measuring, identifying and quantifying the causal effects of policy uncertainty. Recent work is tackling these difficult issues; for example, Baker et al. (2013) construct a news-based index of policy uncertainty and find it helps predicting declines in aggregate output and employment. Our focus and empirical approach are considerably different. We use applied policy and counterfactual policy measures, both of which are observable in our setting, to directly estimate the effects of policy uncertainty on economic activity. In order to identify the effects of TPU we explore both variation over time (capturing the reduction in the probability of a trade war after WTO accession) and across industries (since they would face different tariffs if a trade war broke out).

Our dynamic heterogeneous firms’ model extends Handley and Limão, 2012 (HL) in two important ways. First, firms can invest to both enter foreign markets and to upgrade their technology. This allows the model to account for the fact that China’s exports increased on both the extensive margin (new exporters) and the intensive one (continuing exporters with upgraded technology). The second theoretical contribution is to allow for the exporting country to be large enough to affect the importer’s aggregate outcomes. In particular, we show that reductions in TPU by an importer (e.g. the U.S.) generate an incentive to enter and upgrade by its partner (China) and this leads to a reduction in the importer’s price index. This price reduction is central to the welfare gains from reforms that lower TPU.

The theory is also essential in guiding the estimation and quantification. First, it allows us to construct a theory-consistent measure of uncertainty—the proportion of profits that Chinese exporters would lose if China ever lost its MFN status. Importantly, this pre-WTO measure can be calculated using observable MFN and threat tariffs (so called column 2 tariffs). Second, the model generates a tractable TPU-augmented gravity equation that allows us to consistently aggregate individual firm decisions to the industry level. We then explore the variation in policies and trade to identify the impact of TPU on thousands of products. We find non-parametric and parametric evidence that Chinese export growth in 2000-2005 was higher in industries with higher initial TPU. The policy uncertainty reduction increased Chinese exports to the U.S. by up to 29 log points. That effect falls to 22 log points when we account for general equilibrium price effects. Our approach allows for non-linear effects of uncertainty and controls for applied trade policy barriers, transport costs, unobserved industry heterogeneity and sector specific growth trends.

Understanding the impact of TPU has broad implications. It informs us about the potential impacts of other sources of policy uncertainty, such as U.S. threats to impose tariffs against “currency manipulators”

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6 For evidence on these margins see cf. Amiti and Freund, 2008, and Manova and Zhang, 2009. Other evidence indicates that applied tariff changes can trigger within firm productivity increases (cf. Lileeva and Trefler 2010) so it is plausible that the same may happen due to reductions in TPU. This could account for the evidence of substantial firm-level TFP growth increases in China since 2001 (Brandt et al, 2012).
or revoke unilateral preferences to developing countries. Promoting trade is a central goal of the WTO, but Rose (2004) argues the WTO has not succeeded whereas others argue it has (cf. Subramanian and Wei, 2007). Our work highlights a trade promotion channel that is largely missing from the empirical and theoretical debate on trade agreements, barring recent exceptions discussed below. We also contribute to the long standing question of the aggregate gains from trade. Recent work by Arkolakis et al. (2012) has focused on the costs of applied trade barriers. Our framework highlights the costs of TPU and provides a theory consistent applied policy cost equivalent. We estimate that U.S. TPU reduced its consumers’ real income by at least 0.8 percent each year. This is equivalent to a permanent applied tariff increase of 8 percentage points on Chinese goods.

We also contribute to the literature on trade agreements more broadly. Bagwell and Staiger (1999) argue that the central role of the GATT/WTO agreement is to internalize the terms-of-trade effects imposed by tariffs. There is now evidence that countries possess market power and exploit it when they are not in an agreement but less so after an agreement (Broda et al, 2008; Bagwell and Staiger, 2011; Ludema and Mayda, Forthcoming). Moreover, the welfare cost of trade wars in the absence of such agreements are potentially large—about 3.5% for the U.S. according to some quantitative exercises (Ossa, 2013). But this theory and evidence on the WTO has largely ignored TPU. Recent work by Handley (2014) shows that reducing WTO tariff commitments, and thus the worst case tariffs under the agreement, would increase entry of foreign products. Limão and Maggi (Forthcoming) endogenize policy uncertainty and provide conditions such that there is an uncertainty reducing motive for agreements in a standard general equilibrium model. We contribute to this literature by providing both theoretical and empirical evidence for welfare gains from reducing TPU through trade agreements in a dynamic setting with heterogenous firms.

Our research also complements the recent empirical work on the impact of Chinese exports on developed countries. Bloom et al. (2011) assess the reduced form impact of Chinese exports on wages and employment in the European Union while Pierce and Schott (2012) focus on the U.S. Our paper differs in at least two important ways; first our focus is on the trade and consumer welfare effects, second our structural approach allows us to perform counterfactual exercises. For example, we provide a decomposition of the uncertainty effect and find that a substantial fraction is explained by a mean preserving compression of the tariff, and the rest is due to locking in tariffs below the mean. In the working paper version we also quantify the uncertainty impact of proposed legislation that threatens to impose tariffs of almost 30% on “currency manipulators”.

We present the theory in section 2, followed by the empirical approach, data and estimates in section 3 and conclude in section 4. The appendices contain the proofs as well as details on derivations and empirical implementation.

## 2 Theory

We first describe the basic entry decision problem. We then derive the equilibrium entry decisions for firms, first from the perspective of a small exporting country—one that takes foreign aggregate variables as

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7The latter study appeals to the theoretical TPU mechanism in HL to use U.S. column 2 tariffs as a reduced form determinant of the impact of Chinese exports on U.S. manufacturing employment. However, in HL there is no aggregate impact of TPU on the importer (the European Community) because the exporter is assumed to be small (Portugal). In contrast, the current paper does include an impact of TPU on the importer via the price index and a channel through which a reduced form impact of column 2 tariffs on U.S. outcomes can be justified.

8We find that implementing such legislation in 2012 would have had similar trade effects to removing China’s permanent MFN status in 2005 and a higher welfare cost to U.S. consumers.
given—and then a large one. For ease of exposition we initially focus on export entry decisions in a single industry and then extend them to multiple industries and allow for technology upgrading decisions.

2.1 Demand, Supply and Pricing

Consumers spend a fixed share of their income on a homogeneous good and the remaining on a CES aggregate of differentiated goods, both of which are tradable. Each period consumers observe current economic conditions and choose the optimal quantity of each differentiated good, \( q_v \), to maximize utility subject to their budget constraint. This yields the standard CES aggregate optimal demand \( q_v = EP^{\sigma-1}p_v^\sigma \) where \( \sigma > 1 \) is the constant elasticity of substitution across \( v \) and \( p_v \) is the consumer price. The aggregate demand shifter, \( E \), is the total expenditure in the differentiated sector in that country and \( P = \left[ \int_{v \in \Omega} (p_v)^{1-\sigma} \right]^{1/\sigma} \) is the CES price index over the set of available varieties, \( \Omega \).

The supply side is also standard. There is a single factor—labor—with constant marginal productivity normalized to unity in the homogeneous good; the latter is taken as the numeraire so the equilibrium wage is unity in a diversified equilibrium. In the differentiated sector, there is a continuum of monopolistically competitive firms each producing a variety, \( v \), with heterogeneous productivity \( 1/c_v \). Firms know their time invariant productivity and the distribution of other firms in each market.

The consumer price, \( p_v \), includes an advalorem tariff, \( \tau \geq 1 \), so exporters receive \( p_v/\tau \) per unit (domestic producers face no taxes in their market). The tariff is common to all firms in the differentiated industry. After observing \( \tau \) each firm chooses \( p_v \) to maximize operating profits taking aggregate conditions as given, and correctly anticipating their equilibrium value. We allow for an advalorem export cost, \( d \geq 1 \), so operating profits from exporting are \( (p_v/\tau - dc_v)q_v \). This yields the standard mark-up rule over cost, \( p_v = \tau dc_v\sigma/(\sigma - 1) \), and equilibrium operating profit equal to

\[
\pi(a, c_v) = ac_v^{1-\sigma}
\]

where the relevant economic conditions are summarized by \( a \equiv (\tau \sigma)^{-\sigma} ((\sigma - 1) P/d)^{\sigma-1} E \).

2.2 Policy Uncertainty and Export Entry

The timing and information relevant for export entry are the following. At the start of each period surviving firms observe the state, denoted by \( s \), that includes information about (i) the set of firms active in the previous period; (ii) the realization of the policy and (iii) the values of all model parameters in the start of the period. This information permits each firm to correctly infer market conditions in that state, \( a_s \), and form (rational) expectations about future profits. If entry in a state maximizes the firm’s expected profits net of a sunk entry cost, \( K \), then it will enter and continue to export in the following period with probability \( \beta < 1 \), the exogenous probability of survival. There are no fixed costs and thus no endogenous exit. Since the sunk and marginal costs are known and constant, the only source of uncertainty is the future value of market conditions and the timing of death.

\footnote{To focus on export entry we take as given the mass of differentiated firms operating in their \textit{domestic} market as in Chaney (2008) and provide foundations for this later.}
For any state $s$ the expected value from exporting for any firm $v$ after entry is

$$\Pi_e(a_s, c) = \pi(a_s, c) + E_s \sum_{t=1}^{\infty} \beta^t \pi(a'_s, c)$$

(2)

where we omit the variety subscript for simplicity; $E_s$ denotes the expectation over possible future states conditional on the current state's information set.

If the firm does not expect the state to change, there is no uncertainty about economic conditions and no option value of waiting to enter. In this case the firm will only enter if its cost is below a threshold value, $c^D_s$. We compare later results to this benchmark threshold. To derive it, we equate the present discounted value of profits to the sunk cost:

$$\frac{\pi(a_s, c^D_s)}{1 - \beta} = K \iff c^D_s = \left[ \frac{a_s}{(1 - \beta) K} \right]^{\frac{1}{\sigma - 1}}$$

(3)

If future conditions are uncertain then a non-exporter must decide whether to enter the market today or wait until conditions improve. The optimal entry decision of a given firm in state $s$ maximizes its expected value, given by the Bellman equation:

$$\Pi(a_s, c) = \max \left\{ \Pi_e(a_s, c) - K, \beta E_s \Pi(a'_s, c) \right\}.$$  

(4)

The solution to this optimal stopping problem with aggregate uncertainty will take the form of intervals of $a$ over which a firm will enter. Under reasonable assumptions on the persistence of policy we can show that for sufficiently good economic conditions a firm will enter. Therefore, when $a$ is decreasing in tariffs ($\tau$) the solution is to enter when current tariffs are below a firm specific threshold tariff. Given a continuum of firms distributed over $c$ we have that for any given $a_s$ there is a firm with cost equal to the threshold value, $c^U_s$, given by the entry indifference condition:

$$\Pi(a_s, c^U_s) = \Pi_e(a_s, c^U_s) - K,$$

(5)

and any firms with lower costs will enter in state $s$. To characterize this cutoff and derive testable predictions related to WTO entry we now model the policy.

### 2.3 Policy Regime

The trade policy regime is characterized by a Markov process with time invariant distribution, $\Lambda(\tau_m, \gamma)$, which is conditional on the current tariff, $\tau_m$, and on $\gamma$—a measure of uncertainty defined below. We assume $\tau_m$ can take on one of three values, $m = 0, 1, 2$, ordered in increasing value of $\tau_m$. So $\tau_1$ corresponds to an intermediate value—the temporary MFN status that Chinese producers faced before the WTO—and $\tau_2 > \tau_1$ is the threat value if that status was revoked. We represent the $3 \times 3$ policy transition matrix by

$$M = \begin{bmatrix} \lambda_{22} & \lambda_{21} & \lambda_{20} \\ \lambda_{12} & \lambda_{11} & \lambda_{10} \\ \lambda_{02} & \lambda_{01} & \lambda_{00} \end{bmatrix}$$

(6)

This is an implicit solution for the cutoff if exporters account for a non-negligible fraction of the differentiated sector in the destination market (because $a_s$ will then depend on $c^D_s$).
where $\lambda_{11} = 1 - \gamma$ is the probability of remaining in policy state 1. The probability that temporary MFN status was revoked is $\lambda_{12}$. In the late 1990s and through 2001, China was negotiating accession to the WTO under which it would obtain $\tau_0 \leq \tau_1$ with probability $\lambda_{10}$. We assume that $\Lambda (\tau_{m+1}, \gamma)$ first order stochastically dominates $\Lambda (\tau_m, \gamma)$ for $m = 0, 1$. This condition is useful in characterizing the solution to the firm’s optimal stopping problem as a single threshold (instead of sets of intervals, cf. Dixit and Pindyck, 1994). To simplify the derivation and presentation of the results we assume WTO is an absorbing policy state, so $\lambda_{00} = 1$.

2.4 Small Exporter Equilibrium

Our two objectives in this section are to compare the cutoffs across states and derive how they are affected by changes in policy uncertainty. Tariffs are the only underlying source of policy uncertainty and we initially focus on a small exporting country such that the set of its exporting firms has a negligible effect on the importer country’s aggregate variables. This implies that tariff changes only have a direct impact on market conditions, $a_s$, and will not affect the equilibrium aggregate variables in the importing country, $E$ and $P$.

With a small exporter there is one distinct value of $a_s$ for each $\tau_m$ for any given value of $EP^{\sigma-1}$.

In the proof of proposition 1 (in appendix section A.1) we use this insight and the policy distribution to show that the solution to the Bellman equation in (4) is a single value of economic conditions above which a firm enters. Therefore the indifference condition in (5) will imply one distinct cutoff, $c_{U1}$, for each $\tau_m$. We are interested in the role of uncertainty during the state with intermediate tariffs. In the proof of Proposition 1 we also derive a closed form expression for $c_{U1}$ and show it is proportional to its deterministic counterpart in (3) by an uncertainty factor, $U_1(\omega, \gamma)$.

$$c_{U1} = c_{D1} U_1(\omega, \gamma)$$

$$U_1(\omega, \gamma) \equiv \left[ \frac{1 + u(\gamma) \omega}{1 + u(\gamma)} \right]^{\frac{1}{\sigma-1}}$$

If $U$ is lower than one then there is less entry under uncertainty. To interpret this factor, note that $\omega \equiv \left( \frac{\tau_2}{\tau_1} \right)^{-\sigma} < 1$ is the ratio of operating profits under the worst case scenario relative to MFN (given no other conditions changed). The term $u(\gamma) \equiv \frac{\beta_2 \lambda_{12}}{\beta_1 \lambda_{22}}$ is the average spell that a firm starting at the temporary MFN state expects to spend under column 2.

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Note that we do not need to impose additional general equilibrium structure to solve for these export cutoffs. Uncertainty will change the profits from exporting but these are separable from domestic profits since the wage is pinned down at unity and there is a fixed mass of firms serving their own market.

\[11\] It is also plausible since (i) if China faced column 2 tariffs it would be less likely to transition to the WTO state directly than would be the case if it were in a negotiation/MFN stage, i.e. $\lambda_{20} \leq \lambda_{10}$, and (ii) it is less likely to transition from WTO to column 2 than before, i.e. $\lambda_{02} \leq \lambda_{12}$.

\[12\] In practice the WTO does not end all TPU but the evidence we will consider suggests that it did end TPU regarding column 2 tariffs.

\[13\] Note that we do not need to impose additional general equilibrium structure to solve for these export cutoffs. Uncertainty will change the profits from exporting but these are separable from domestic profits since the wage is pinned down at unity and there is a fixed mass of firms serving their own market.
Proposition 1 uses these definitions to summarize the effects of TPU on the export entry cost cutoffs when foreign exporters have a negligible impact on the importer’s aggregate variables.

**Proposition 1: Policy Uncertainty and Export Entry (small country).**

Under a regime $\Lambda(\tau, \gamma)$ with policy uncertainty and where tariff increases are possible, the entry cutoff under the temporary MFN state, $c_U^1$, is

(a) unique and proportional to its deterministic counterpart, $c_D^1$, by the uncertainty factor, $U_1(\omega, \gamma)$.

(b) lower than its deterministic counterpart ($c_U^1 < c_D^1$) and decreasing in policy uncertainty ($dc_U^1/d\gamma = dU_1/d\gamma < 0$ for all $\gamma$)

(c) lower than the agreement cutoff ($c_U^1 < c_D^0 = c_D^0$).

**Proof:** See Appendix A.1.

Part (a) summarizes the cutoff relationship in (7). To show part (b) we derive necessary and sufficient conditions for $U_1 < 1$. If we evaluate (8) at either $u(\gamma) = 0$ or $\omega = 1$ we verify that it is unity so it is necessary to have the possibility of tariff increases for $c_U^1 < c_D^1$. Thus our subsequent estimation can nest the special case where firms never expect column 2 tariffs. Evaluating (8) at $\omega < 1$ and $u(\gamma) > 0$ we find $U_1 < 1$, so the condition is sufficient. While TPU can lead to lower or higher tariffs, it is only the possibility of high tariffs that affects entry; if the MFN state is temporary ($\gamma > 0$) but tariff increases are not possible ($\lambda_2 = 0$) then uncertainty has no impact on entry. If tariff increases are possible, then entry is reduced. Moreover, this effect is monotonic ($dc_U^1/d\gamma < 0$ for all $\gamma$) so even before an agreement (i.e. in the temporary MFN state) we can test if there are significant changes in policy uncertainty, e.g. in years where an MFN vote was more likely. Part (c) compares the cutoff before and after WTO entry. If the entry only eliminated uncertainty then the effect would be the same as in part (b). If there is also a reduction in applied tariffs there will be additional entry; it is important to control for any such reductions.

In sum, under the conditions in proposition 1, the set of exporters is higher under the agreement than under temporary MFN. The explicit solution for the entry cutoff allows us to derive its elasticity with respect to $\gamma$, which is larger for industries with higher potential losses under the worst case scenario.

### 2.5 Decomposition of the impacts of uncertainty shocks

The theoretical results focus on a particular measure of policy uncertainty, the probability that the current policy will change. In the empirical section we will decompose the effects of shocks to this measure of uncertainty into two effects. Consider starting at the temporary MFN state with $\tau_1$ and examining the effect of a shock that eliminates uncertainty but does not change the applied tariff. If $\tau_1$ was at the long-run mean of original tariff process then this uncertainty reduction is exactly a mean preserving compression of tariffs, or a pure risk reduction. However, if $\tau_1$ was below its long-run mean, as will be the case in our application, then the reduction in $\gamma$ has the additional effect of locking in tariffs below their expected value under uncertainty.

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14 Proposition 1 applies the same basic insight in Handley and Limão (2012) to a policy process with state dependence even after the policy shock.

15 In the Online Appendix C.2 we also prove that under the conditions in the proposition we have less entry under column 2, i.e. $c_U^2 = c_D^2 < c_U^1 < c_D^1$. The reason why $c_U^1 = c_D^2$ is that there is no threat that conditions will worsen.

16 It is straightforward to show in this three state process that when state 1 has a policy $\tau_1$ equal to the long-run mean then a decrease in $\gamma$ induces a mean preserving compression of the initial conditional policy distribution, $\Lambda(\tau_1, \gamma)$.
2.6 Aggregate Effects

2.6.1 Setup and equilibrium definition

We now examine policy uncertainty when differentiated goods’ exports are sufficiently large to affect the destination market’s aggregate expenditure, \( E_s \), and price index, \( P_s \). To do so and close the model in a tractable way we make the following additional assumptions.

A1. There is no borrowing technology available across periods. This implies current expenditures must equal current income each period for each individual.

A2. All individuals have labor endowments of \( \ell \) each period. A fraction of individuals are workers and the rest are entrepreneurs. The latter have constant mass \( N \) and are endowed with a blueprint for a variety that is embodied in the marginal cost parameter \( c_v \). Entrepreneurs receive the profits of their respective variety and any import policy revenue, which is rebated lump-sum.

A3. The constant expenditure share of per period utility on differentiated goods is \( \mu > 0 \) for workers and zero for entrepreneurs.

A4. The two countries have the same preference structure.

We highlight five implications of this structure. (1) A2 implies that the only source of worker income is the wage, which is still pinned down by the marginal product of labor in the numeraire—unity.\(^{17}\) (2) The constant equilibrium wage and A1-A3 together imply that expenditure on differentiated goods is constant: \( E = \mu L \) where \( L \) is the mass of workers; so the price index is the only aggregate endogenous variable that is uncertain in each country. (3) The indirect utility for workers is \( \tilde{\mu}P_s - \mu_s \) in each state.\(^{18}\) (4) A3 implies that entrepreneurs have linear utility so the export entry decision of a risk neutral entrepreneur that survives each period with probability \( \beta \) is obtained by solving the Bellman equation in \( 4 \).\(^{19}\) (5) A4 rules out third country effects.

The price index for differentiated goods in state \( s \) depends on the union of imported and home varieties, \( \Omega_s = \Omega^x_s \cup \Omega^h \):

\[
P_s^{1-\sigma} = \int_{\Omega^x_s=\Omega^y_s \cup \Omega^h} (p_{vs})^{1-\sigma} \, dv = \int_{\Omega^x_s} (\tau_m c_v / \rho)^{1-\sigma} \, dv + \int_{\Omega^h} (c_v / \rho)^{1-\sigma} \, dv
\]

where \( \rho \equiv \frac{\sigma-1}{\sigma} \).

Before deciding to export, firms form rational expectations about the expected price index, \( P^e_s \). In equilibrium \( P^e_s = P_s \) given the following information structure. At the start of each period \( t \) a surviving firm knows its cost, \( c_v \), and there is a common knowledge information set, denoted \( i_s \), that includes (i) the fixed exogenous parameters of the model including the survival rate, \( \beta \), and the time invariant set of potential varieties in each country, \( \Omega \); (ii) the structure of the model including the entry decision rule; (iii) the current realization of the policy, and; (iv) the equilibrium set of varieties sold in each market in the previous period, denoted by \( \Omega_{t-1} \). The state, \( s \), is defined by the combination of the realized policy at \( t \) and \( \Omega_{t-1} \).

\(^{17}\)This follows because the population in each country is sufficiently large for the numeraire to be produced in equilibrium.

\(^{18}\)The constant is \( \tilde{\mu} \equiv w\ell \mu^\mu (1-\mu)^{(1-\mu)} \) and \( w = 1 \).

\(^{19}\)We rule out the possibility that entrepreneurs are credit constrained by assuming that their endowment \( \ell \geq K \), so they can always self-finance the sunk cost in a single period even if it exceeds that period’s operating profits.
We define the equilibrium as the following set of prices and quantities in each country and state $s$: (a) a demand vector for the differentiated and numeraire good, $q_s$; (b) a market entry decision for each differentiated firm $v$ and a distribution of active firms, $\Omega_s$, with prices, $p_s$; (c) an expected and actual price index, $P^e_s$ and $P_s$, and; (d) labor demands for the differentiated and numeraire goods and a wage, that satisfy the following conditions: (i) the numeraire good market clears; (ii) workers maximize utility subject to their budget constraint taking their factor endowments and all prices as given; (iii) entrepreneurs maximize utility subject to their budget constraint taking as given their factor endowment, technology, wage, $P^e_s$, the policy regime ($\Lambda$) and its lump-sum revenue, and all other information in $i_s$; (iv) $P_s = P^e_s$ due to rational expectations (see Online Appendix C.3), and (v) the labor market clears.

Since TPU now also affects the price index there will be aggregate uncertainty, transition dynamics, and a larger state space given the many possible distributions of active firms. We derive the deterministic policy equilibrium to show how the transition dynamics operate and then the equilibrium under uncertainty.

### 2.6.2 Deterministic policy equilibrium

In our setting each firm’s decision to start exporting is independent of its home market economic conditions. Thus we can solve for the cutoff and price index for each country separately. If the policy is expected to be permanently fixed at some value $\tau_m$ then the only exporters to this market are the firms with cost below the deterministic cutoff, $c^D_s(P^D_s, \tau_m)$, given by (3). To obtain the price index we evaluate (9) using the set of available foreign varieties, which is a function of the cutoff, $\Omega_s(c^D_s)$. So we can express the deterministic price index as $P^D_s(c^D_s, \tau_m)$.

This pair of equations for each market has a unique solution. For any given fixed tariff value the entry schedule, $c^D_s$, is linear and increasing in $P^D_s$ and $c^D_s|_{P_s \rightarrow 0} = 0$ whereas $P^D_s$ is positive and decreasing in $c^D_s$.

Therefore these schedules intersect only once at a point such as $I$ in Figure 2.

### 2.6.3 Unanticipated shocks and transition dynamics

We now show that the adjustment process to an unanticipated permanent change in the policy depends on the direction of that change and derive its impacts on different outcomes.

Figure 2 depicts an unexpected permanent decrease in the tariff starting from the initial equilibrium point, denoted by $I$. At any given value of the price index, the tariff reduction increases export profits and thus generates entry. For a given price index the new cutoff would be at $PE$ but if the exporting country is large then $P^D_s$ pivots down—since for any cutoff the consumer prices are lower—and the new equilibrium is at $GE$. We will show that $P$ falls and counteracts some but not all of the direct effect of the tariff.

The steady state effect of an unexpected permanent tariff change is obtained by totally differentiating $c^D(P^D_s, \tau_m)$ in (3) and $P^D(c^D_s, \tau_m)$, which yields

$$
\hat{c}^D = \frac{\varepsilon_{c} - 1/\rho}{1 - \varepsilon_c} \hat{\tau},
$$

$$
\hat{P}^D = \frac{\varepsilon_{c} - 1/\rho}{1 - \varepsilon_c} \hat{\tau}.
$$

\[20\] The only domestic condition that is relevant is the wage but this is constant and fixed over time by its marginal product in the numeraire sector.

\[21\] To be clear, $P^D(c^D_s, \tau_m) = \left[ N \int_0^{c^D_s} (\tau_m c_v/\rho)^{1-\sigma} dG(c) + \int_{\Omega_s} (c_v/\rho)^{1-\sigma} dv \right]^{1/(1-\sigma)}$

\[22\] A similar figure applies to the other country, but the equilibrium levels may be different depending on its tariff, domestic mass of producers and productivity distribution.
where \( \dot{x} = d \ln x, \varepsilon_x = \frac{\partial \ln p(\tau, c)}{\partial \ln x} \) evaluated at the original tariff. The cutoff increases if \( \hat{\tau} < 0 \) since \( \varepsilon_c \leq 0 \) and \( \varepsilon_x \leq 1 < 1/p \). The higher cutoff due to the tariff reduction has two implications. First, even in general equilibrium firms benefit from facing lower tariffs in their industry in the export market. Second, when the tariff falls there are no transition dynamics because entry allows the economy to move immediately from \( I \) to \( GE \). The consumption gain from trade liberalization that accrues to workers is \(-\mu \hat{P}\), the proportional change in the price index weighted by the differentiated goods’ share in expenditure. For domestic firms an unanticipated unilateral tariff reduction that decreases \( P \) reduces demand for their varieties and thus their domestic profits by \( \hat{\pi}_v = (\sigma - 1) \hat{P} \), which is similar across all domestic \( v \).

Now consider an unanticipated permanent increase in the tariff. The steady state values are obtained as before and shown in Figure 3 as \( 1^D \) — the initial equilibrium — and \( 2^D \) — the final one. However, now there are transition dynamics because after a tariff increase firms with costs above the cutoff will continue to export until they die. A general state after a permanent tariff increase is denoted by \( s = 2T \), which identifies the tariff state, \( \tau_2 \), and the period \( T = 0, 1, \ldots \) measuring the time since the tariff increase.

The transition is described by a jump from \( 1^D \) to a point on the entry schedule between \( 2^{TR} \) and \( 2^D \) followed by an adjustment over time to \( 2^D \). We show this in two steps. First, for a given entry schedule evaluated at the higher tariff, \( c^D_{2T} \), the price index first jumps, due to the direct effect of higher tariffs, and then increases monotonically to its steady state value, \( P^D_2 \). If there was no death shock then the economy would immediately adjust to \( P^{TR} \) in Figure 3 and stay there. This price is higher than the initial equilibrium (because of the higher tariff) but lower than \( P^D_2 \) — the steady state when some firms exit. The monotonic increase of the price index towards its steady state, \( P^D_2 \), is driven by the exogenous death shock process. Each period following the tariff increase, a fraction \( 1 - \beta \) of firms die and a similar fraction are born, but new firms with cost above the new cutoff do not enter the export market. So, in the period when the tariff increases we jump to a price higher than \( P^{TR} \), i.e. a point above \( 2^{TR} \).

After the initial jump the adjustment occurs along the deterministic entry schedule from \( 2^{TR} \) to \( 2^D \) such that \( c^D_{2T} = c^D (P^D_{2T}, \tau_2) \). During the transition new firms still face an optimal export entry decision problem since there is uncertainty about the timing of death. If a firm enters today it pays \( K \) and obtains \( \sum_{t=1}^{\infty} \beta^{t-T} \pi(a_{2t}, c) \) and if it waits then its payoff from export is zero today. But if it is just indifferent between the two then in the following period it will surely enter and obtain a PDV of \( \sum_{t=1}^{\infty} \beta^{t-1} \pi(a_{2t}, c) \) because, after the shock, aggregate conditions will be improving (recall that \( P \) increases as firms exit). Therefore the marginal entrant at \( T \) is the firm with cost equal to the cutoff defined by \( \pi(a_{2T}, c = c^D_{2T}) = (1 - \beta) K \). For this firm the extra profit flow from entering today relative to tomorrow, \( \pi(a_{2T}, c) \), is just enough to cover the extra cost paid today instead of next period, \((1 - \beta) K \). Therefore the entry schedule has the same functional form as the one derived for the deterministic steady state. The equilibrium cutoff value in transition, \( c^D_{2T} \),

---

23. We have \( \varepsilon_x \leq 1 \) since the highest possible elasticity would occur if all goods (including domestic) were taxed at \( \tau \) and the partial elasticity of \( P \) with respect to it would then be \( 1 \).

24. The result is not surprising in the single industry case but not obvious in the multi-industry setting. In our working paper we provide a sufficient condition to show it extends to a multi industry setting, namely if the proportional tariff change is identical across industries. We also provide a weaker sufficient condition for the ranking of economic conditions when the tariff liberalization is not similar across industries.

25. We will derive an analogous expression under uncertainty and show how to estimate \( \varepsilon_x \) and \( \varepsilon_c \) under specific productivity distributions to calculate the impact of changes in TPU.

26. Thus the model naturally generates variation in exporter status noted in different data sets for firms of similar productivity during transition: some firms do not export even though they have higher productivity than legacy firms.
can be related to the “steady state” equilibrium, $c^D_2$, as follows

$$
c^D_{2T} = \left[ \frac{a_{2T}}{(1-\beta)K} \right]^{\frac{1}{\gamma}} = c^D_2 \left[ \frac{a_{2T}}{a_2} \right]^{\frac{1}{\gamma}}
$$

(11)

Note that $a_{2T}/a_2$ is equal to the ratio of profits at $T$ relative to steady state under $\tau_2$. It is lower than unity as long as the price index at $T$, is below its steady state, $P^D_2$, as we formally show in Appendix C.3.

### 2.6.4 Entry, prices and welfare under policy uncertainty

We now build on the deterministic analysis to examine TPU with large countries. After a negative shock there is sluggish net exit; conditions for potential entrants are worse in transition than in steady state, which we need to take into account to compute the value functions.

**Entry and Prices**

To derive the stochastic process for the economic conditions, $a_s$, we first describe the state space. A state is any distinct combination of the current tariff realization and the distribution of active firms in the previous period. In the deterministic setting the possible states were $s = 0, 1, 2T$ where $T \geq 0$ is the number of periods since the tariff increase from $\tau_1$ to $\tau_2$. The ranking of current economic conditions is $a_0 > a_1 > a_{2T}$ and $a_{2T} > a_{2T-1}$ because the price index is increasing over time under $\tau_2$; so the same ranking will hold under uncertainty, at least around the deterministic outcome.

Under uncertainty, a tariff increase to $\tau_2$ also generates transition dynamics. But now the number of states under $\tau_2, 2T$, is expected to be finite because there is a constant probability of returning to $\tau_1$. Our setting simplifies the potentially large state space because if the policy is at the MFN or WTO value there is no possible history of better conditions and no history dependence. This implies that the states under $\tau_{m=0,1}$ are fully described by their policy realization.\footnote{Tariffs are lowest at the WTO state so no better conditions are possible. The reason why there is no history of better conditions at $s = 1$ is that the only other state that would yield such conditions is the WTO, which we assume is absorbing.}

In sum, the state space is finite and $a_s$ follows a Markov process with time invariant distribution, $\tilde{\Lambda}(a_s, \gamma)$. Any positive transition probability from $s$ to $s'$ in $\tilde{\Lambda}(a_s, \gamma)$ is given by the policy transition probability and in the proof of proposition 2 we show that the properties we derive for $\tilde{\Lambda}(a_s, \gamma)$ are sufficient to ensure that a threshold rule for entry also applies here. Namely, a firm with cost $c_v$ enters in state $s$ if economic conditions are above a unique threshold $a^U_s(\gamma, c_v)$. Since the general approach to determine these cutoffs is similar to section (2.2) we describe the main results below and provide the details in Appendix A.2.

Under the WTO, the entry cutoff value is still given by (3) evaluated at $\tau_0$ and taking into account the price index in (9). The worst case entry schedule is still given by (11). The reason is that a firm that is indifferent between entering $T$ periods after transitioning to $\tau_2$ or waiting will surely enter at $T+1$ if it survives because either tariffs fall or aggregate conditions improve as other firms exit. Thus a firm is indifferent about entering in a state $2T$ if $\Pi_e(a_{2T}, c) - K = \Pi(a_{2T}, c)$, which yields exactly the entry schedule in (11). The transition dynamics will also be similar to what we derived under the unanticipated permanent tariff increase: a jump in $P$ followed by exit and a further increase in $P$ along the transition path. The main difference is that under uncertainty we start at a different equilibrium value for $P_1$ and $c^U_1$ under the MFN state.
The MFN equilibrium is now at a point such as $1_U$ in Figure 3. The price index schedule is the same as in the deterministic case since it is independent of $\gamma$ for a given cutoff value. Moreover, we already established that the deterministic equilibrium is at $1^D$. Thus we need to show that uncertainty in this state entails a lower entry cutoff by solving the Bellman equation in (4). The key difference relative to proposition 1 is the price index effect that now requires us to derive a new stochastic process for $a_s$, as described above, and take transition dynamics into account when a bad shock arrives. We provide a solution method in Appendix A.2 and employ the indifference condition in (5) to derive the following cutoff in the temporary MFN state:

$$c_1^U = \left[ \frac{a_1}{(1-\beta)K} \right]^{\frac{1}{\sigma-1}} U_1(\tilde{\omega}, \gamma) = a^D U_1(\tilde{\omega}, \gamma) \frac{P_1}{P^D_1}$$  \hspace{1cm} (12)

$$U_1(\tilde{\omega}, \gamma) = \left[ \frac{1 + u(\gamma)\tilde{\omega}}{1 + u(\gamma)} \right]^{\frac{1}{\sigma-1}}$$  \hspace{1cm} (13)

The difference in the uncertainty factor, $U_1$, relative to (8) for the baseline case is the term $\tilde{\omega} = (\tau_2/\tau_1)$, reflecting the ratio of the average operating profits under the worst case scenario relative to state 1. But now it also reflects a general equilibrium effect given by

$$g = (1-\beta\lambda_{22}) \sum_{T=0}^{\infty} (\beta\lambda_{22})^T \left( \frac{P_{2T}}{P_1} \right)^{\sigma-1} \geq 1$$  \hspace{1cm} (14)

capturing average economic conditions (other than tariffs) after a transition from MFN to column 2. This dampens the direct uncertainty effect because the increased price index under column 2 generates higher profits for the surviving firms. In addition, the price index is higher under uncertainty, $P_1 > P^D_1$, because there is less entry, which further offsets the direct effect of uncertainty on entry in general equilibrium. Proposition 2 establishes the effects of TPU in with general equilibrium price effects between large countries.

**Proposition 2: Policy Uncertainty, Export Entry and Prices.**
Under a regime $\Lambda (\tau_m, \gamma)$ with policy uncertainty firms enter if $c_v \leq c^U_s$. Moreover, when tariff increases are possible

(a) the unique cutoff under the MFN state is $c_1^U = c^D U_1(\tilde{\omega}, \gamma) P_1/P^D_1$ where $U_1(\tilde{\omega}, \gamma)$ is in eq. (13).

(b) eliminating MFN uncertainty ($\gamma = 0$) increases entry and decreases the price index at MFN: $c_1^U < c^D_1$ and $P_1 > P^D_1$;

(c) the trade agreement increases entry and decreases the price index: $c_1^U < c^U_0 = c^D_0$ and $P_1 > P_0 = P^D_0$.

Proof: See Appendix A.2.

There are several differences relative to Proposition 1. Part (a) summarizes the relationship between the cutoff under uncertainty and the deterministic cutoff. Relative to Proposition 1 there is now a smaller effect of policy uncertainty on entry because of the price effects. Nevertheless, part (b) establishes that eliminating policy uncertainty will nonetheless increase entry in the MFN state. To see this consider point $1^U$ in Figure 3 as the MFN equilibrium and then setting $\gamma = 0$ so $U = 1$. Doing so implies that, at the original price index level, the cutoff increases, so the equilibrium price moves down along the original schedule.

Part (c) compares the temporary MFN outcome with the agreement. If the tariffs under MFN and the agreement are the same, then we can apply part (b). If they are lower under the agreement we have additional entry and a lower price. This relationship between parts (b) and (c) is useful because it suggests that if we
Empirically control for applied barrier changes then we can use a switch to the agreement state to infer what would have happened if only uncertainty was eliminated.

The empirical analysis explores variation in tariffs across industries. Therefore in the working paper we provide sufficient conditions for parts (a), (b) and (c) to hold in a multi-industry setting.

**Consumer welfare**

We now derive the effect of policy uncertainty changes on consumer welfare. To quantify these effects empirically we focus on impacts around the deterministic equilibrium, which we showed exists, is unique, and a special case of the stochastic model when $\gamma = 0$.

For given values of $\bar{\omega}$ the only direct impact of $\gamma$ occurs through $U_1(\bar{\omega}, \gamma)$ via changes in the transition probabilities. Differentiating (13) we obtain

$$
\gamma \times \left. \frac{d \ln U_1(\bar{\omega}, \gamma)}{d \gamma} \right|_{\gamma=0} = \left. \frac{d \ln \hat{\omega}_{\gamma=0} - 1}{\sigma - 1} \right|_{\gamma=0} \tag{15}
$$

where we obtain $\hat{\omega} |_{\gamma=0}$ by evaluating $g$ at $P_1^D$ and $P_2^{D_T}$. The effect of $\gamma$ on $U$ will then affect the cutoffs, which will in turn impact the price index and consumer welfare. The price index in the MFN state under uncertainty can be written as a function of the tariffs and cutoffs in that state, $P_1(\tau_1, c_1)$. From (12) we know that $c_1^U(U_1, P_1, \tau_1)$ is log linear in its arguments. The impact of changing $\gamma$ is found by replacing (15) in the following system and solving it:

$$
\left. \frac{d \ln c_1^U}{d \gamma} \right|_{\gamma=0} = \left. \frac{d \ln U_1}{d \gamma} \right|_{\gamma=0} + \left. \frac{d \ln P_1}{d \gamma} \right|_{\gamma=0} \tag{16}
$$

$$
\left. \frac{d \ln P_1}{d \gamma} \right|_{\gamma=0} = \varepsilon_c \left. \frac{d \ln c_1^U}{d \gamma} \right|_{\gamma=0} \tag{17}
$$

where $\varepsilon_c \equiv \frac{\partial \ln P(c_1)}{\partial \ln c_1} |_{\gamma=0}$. We solve the system and verify that uncertainty increases the price index

$$
\gamma \times \left. \frac{d \ln P_1}{d \gamma} \right|_{\gamma=0} = u(\gamma) \frac{\hat{\omega}_{\gamma=0} - 1}{\sigma - 1} \varepsilon_c > 0 \tag{18}
$$

where $\varepsilon_c < 0$ due to love of variety. Uncertainty increases the price index if $\hat{\omega} < 1$ as Proposition 2 shows.

Proposition 3 summarizes the impact of policy uncertainty on $P$ and consumer welfare.

**Proposition 3: Policy Uncertainty and Consumer Welfare.**

If tariff increases are possible in the MFN state then

(a) eliminating policy uncertainty increases consumer welfare in the MFN state.

(b) expected consumer welfare is lower than under an agreement that eliminates policy uncertainty.

(c) the expected consumer welfare from eliminating uncertainty in the MFN state is given by (20)

Proof: See Appendix A.3.

---

28 Changes in $\gamma$ affect current conditions and therefore the future price path reflected in $\omega$. However, the latter are indirect general equilibrium effects and are multiplied by $\gamma$ and so they are negligible when evaluating around $\gamma = 0$.

29 The expression in (18) holds for general distributions of productivity; in Online Appendix C.1.2 we show what it implies under the standard Pareto distribution that we employ in the estimation.
To understand part (a) recall that the price index in the MFN state is higher under uncertainty (proposition 2) and the period indirect utility of consumers in state $s$ is $\hat{\mu}_P^{-\mu}$. Therefore a reduction in uncertainty increases consumer welfare in the MFN state. Part (b) focuses on the expected welfare for consumers, which, starting at a state $s$, is $W_s = \hat{\mu}_P^{-\mu} + \beta \mathbb{E}_s W'_{s}$ for $s = 0, 1, 2T$—the sum of current utility and the expected continuation value. We prove that starting at policy state $m = 1$ expected welfare is

$$W_1 (\gamma) = \sum_{m=0,1,2} n_m (\gamma) w_m (\gamma)$$

(19)

where $n_m$ is the number of periods a consumer expects to spend in each policy state and $w_m (\gamma) \equiv \mathbb{E}_{s \in m} \hat{\mu}_P^{-\mu}$ is the average period utility associated with each policy state.\footnote{We have $w_m = \mu_\mu^{-\mu}$ and $w_2 = (1 - \beta \lambda_{22}) \sum_{T} (\beta \lambda_{22})^T \mu_\mu^{-\mu}$. We derive $n_m$ and other properties in Online Appendix C.5.} Recalling that $P_0 < P_1 < P_2$ (Proposition 2) we then have $w_0 > w_1 > w_2$ so expected welfare is also highest under the agreement state (the state with most weight on $w_0$).

In the proof of part (c) we show that using (19) to derive the first order effects of eliminating uncertainty on expected welfare at the MFN state we obtain

$$\ln \frac{W_1 (\gamma > 0)}{W_1 (\gamma = 0)} \approx -\mu \gamma \times \frac{d\ln P_1}{d\gamma} |_{\gamma=0} + \sum_{m=0,2} \frac{\beta \lambda_{1m}}{1 - \beta \lambda_{mm}} \ln \mathbb{E}_{s \in m} \left( \frac{P^D}{P^D_{1}} \right)^{-\mu}$$

(20)

Changes in uncertainty while at the MFN state have two effects on expected welfare: they affect the average utility in each policy state and its expected duration. Around the deterministic equilibrium these effects reduce respectively to the first and second term on the RHS of (20). The first is the increase in the price index in the MFN state, which is negative and given by (18). The second is the increased probability of switching out of MFN; the latter effect is positive if, as is the case in our application, the MFN tariff is not too close to the column 2 tariff.\footnote{In the empirical section we quantify these effects and show how they can be extended to provide (i) a tariff cost equivalent of policy uncertainty and (ii) a decomposition into a pure risk and expected mean effect.}

**Domestic profits and employment in differentiated sector**

A reduction in policy uncertainty lowers the domestic profits of firms in the import market. In our setting this reduction occurs via the price index alone and is equal to $d\ln \pi (a_1, c_v) / d\gamma = (\sigma - 1) \times d\ln P_1 / d\gamma$. When all firms are active in their domestic market this is also the aggregate effect on domestic profits and quantities. These effects are independent of whether the other country changes its policy uncertainty.\footnote{If China also reduced its TPU there would be an offsetting effect on profits for new U.S. exporters to that market.} If we identify the differentiated sector with manufacturing then the proportional change in employment in that sector would be given by this same expression. This implies that a reduction in TPU reduces employment in the manufacturing sector of the importing sector. In this setting the employment effects of TPU have no additional welfare consequences for workers because they can move costlessly to the numeraire sector and receive the same wage.

### 2.6.5 Technology Upgrading and Industry Aggregation

We now examine how changes in policy uncertainty translate into export growth. Since the main goal is to derive a tractable gravity equation for structural estimation we extend the baseline model along two
dimensions. First, we model an impact of uncertainty on the intensive margin of firm’s exports, which accounted for a substantial fraction of China’s export growth. Second, we allow for industry variation in both technology parameters and policies, which is necessary for our identification strategy.

**Technology Upgrade**

Products already exported from China to the U.S. in 2000 account for 85% of export growth in 2000-2005, even at a very disaggregated level (HS-10). Thus for TPU to have a significant impact on export growth there has to be a plausible intensive margin channel. The channel we model here is irreversible investments by incumbent exporters to upgrade their technologies, which is consistent with the large increases in TFP growth of Chinese firms since WTO accession.

To illustrate the main points in the simplest setting consider upgrades that are specific to an export market. In particular, suppose that exporters can incur an additional sunk cost, \( K_z \), to reduce the marginal export cost to a fraction \( z < 1 \) of the baseline cost \( d \). The operating profits are then \( \pi_v = a_s (zc_v)^{1-\sigma} \).

In the working paper we show that the upgrading decision is similar to the entry decision in that it also takes the form of a cutoff cost, \( c_{sz}^U = \phi c_{s}^U \). The upgrade cutoff is proportional to the entry cutoff by a constant upgrading parameter, \( \phi \equiv \left((z^{1-\sigma} - 1) \frac{K}{K_z}\right) \frac{1}{\sigma} \). The upgrade cutoff is lower than the entry one if the marginal benefit from upgrading is sufficiently high relative to its sunk cost. This implies that the marginal entrant does not upgrade. The entry cutoff solutions will be similar to those we derived, but only the more productive exporters will upgrade. Since \( \phi \) is independent of tariffs the elasticity of the upgrade cutoff with respect to tariffs and uncertainty is the same as the elasticity of the entry cutoff—a result we use to consistently aggregate over all firms in an industry regardless of upgrade status.

**Multi-industry Aggregation**

We define an industry \( V \) as the set of firms that draw their productivity from a similar distribution, \( G_V (c) \), and face similar trade barriers. The basic structure of the model is otherwise unchanged. Namely, the policy regime is still described by a Markov process, \( \Lambda (\tau_{mV}, \gamma) \) with \( m = 0, 1, 2 \) and it applies to each \( V \), so it appropriately captures the fact that in our empirical setting if any industry \( V \) moved from the MFN status to the WTO (or column 2) then all industries would face the same policy state.

The endogenous set of exporters in a given state and industry is now denoted by \( \Omega_{sV} \) and their union is \( \Omega \)—the aggregate set used in our earlier definitions of the price index. Specifically, we denote \( \Omega_{VC} \) as the set of varieties in industry \( V \) produced by firms in country \( C = \text{ch(ina)}, \text{o(ther)} \) so \( \Omega = \cup_{V,C} \Omega_{VC} \) and the price index, \( P \) can then be written as

\[
P^{1-\sigma} = \int_{v \in \Omega} (p_e)^{1-\sigma} dv = \sum_{V,C} \int_{v \in \Omega_{VC}} (p_e)^{1-\sigma} dv + \sum_{V,C \neq ch} \int_{v \in \Omega_{VC}} (p_e)^{1-\sigma} dv \quad (21)
\]

---

33 We are not aware of any direct evidence of the impact of foreign tariffs on Chinese productivity but Brandt et al (2012) find that firm-level TFP growth in manufacturing between 2001-2007 is about three times higher than prior to WTO accession, 1998-2001. Moreover, the TFP growth in the WTO period is higher for larger firms, which is consistent with our model’s prediction that those are the most likely to upgrade.

34 An interpretation of the advalorem export cost is that it represents a portion of the export specific freight, insurance, labeling, or cost of meeting a product standard. The firm can invest in a lower marginal cost technology to achieve these. Alternatively, a firm has a plant that produces only for exporting and it invests in production technology that is specific to that plant.

35 Naturally, each industry could go to different tariff levels, as long as their ranking across policy states is identical.
The deterministic policy entry cutoff is still defined by the expression in [3]. However, this is now an implicit solution since \( P \) depends on all industry cutoffs. The elasticity of entry with respect to tariffs now requires comparative statics on a system of equations that determine the foreign exporter entry cutoffs in each of the \( V \) industries and one equation for the domestic price index.\(^{36}\)

The revenue for a given firm from exporting in state \( s \) is \( p_{sv}q_{sv}/\tau_{sv} \). When we use the optimal price and quantity previously derived and assume export trade costs unrelated to policy that are similar across firms in the same industry \( V \) we obtain firm sales equal to \( a_{sv}\sigma (c_{sv})^{1-\sigma} \) for a non-upgrader and \( a_{sv}\sigma (z_{sv}c_{sv})^{1-\sigma} \) for upgraders. The economic conditions variable, \( a_{sv} \), still depends on aggregate income and price index but, now allows for industry specific trade costs (tariffs and exogenous export costs). In our working paper we show that proposition 1, 2 and 3 can be extended to this multi-industry setting and yield similar cutoff expressions, which now depend on industry specific trade costs.\(^{37}\)

The empirics focuses on a switch from the MFN to the WTO state, so our derivation below focuses on these. The mass of exporting firms in these states is equal to \( N_V \times G_v(e_V^U) \), the total number of producers in industry \( V \) in the export country times the fraction of these with costs below the cutoff (since \( G_v \) is the CDF of productivity). Thus exports in industry \( V \) aggregates over the set of firms that upgrade and those that do not as follows.\(^{38}\)

\[
R_sV = a_{sv}dV^{1-\sigma}\sigma N_V \left[ \int_0^{\phi_V e_V^U} (z_Vc)^{1-\sigma} dG_V(c) + \int_{\phi_V e_V^U}^{e_V^U} c^{1-\sigma} dG_V(c) \right] \quad \text{for } s = 0, 1. \quad (22)
\]

We assume that productivity in each industry is drawn from a Pareto distribution bounded below at \( 1/c_V \), so \( G_v(c) = (c/c_v)^k \) and \( k > \sigma - 1 \). Under this assumption we can obtain sharper predictions, nest a standard gravity model in our framework, and provide precise conditions under which we can identify the impact of uncertainty on exports. We integrate the cost terms in \( (22) \), use the definition of \( a_{sv} \), and \( e_V^U \), and take logs to obtain an uncertainty augmented gravity equation in terms of structural parameters,

\[
\ln R_{sv} = (k - \sigma + 1) \ln U_s (\tilde{\omega}_V, \gamma) - \frac{k\sigma}{\sigma - 1} \ln \tau_{sv} - k \ln d_V + k \ln P_s + \frac{k}{\sigma - 1} \ln \mu L + \ln \zeta_V + \ln \tilde{\omega}_V. \quad (23)
\]

The equilibrium expenditure on all differentiated goods is captured by \( \mu L \), the terms \( \tilde{\omega}_V \) and \( \zeta_V \) are combinations of industry specific parameters that are time invariant.\(^{39}\) Without either policy uncertainty, \( U_s = 1 \), or upgrading, \( \zeta_V = 1 \), \( (23) \) reduces to a standard gravity equation (cf. Chaney, 2008). Finally, all else equal, upgrading increases export levels but not the "elasticity" of industry exports with respect to \( U_s \). Thus we can aggregate sales from all firms to estimate the impact of uncertainty on their industry exports without requiring additional information on which firms upgrade.

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\(^{36}\)We provide details for the implicit solution in Appendix A.4 and derive the elasticities in the Online Appendix C.1.

\(^{37}\)Proposition 1 extends without qualification. A simple way to see why the results in proposition 2 can also be extended is to consider a special case where all industries have a similar proportional change in tariffs so \( \omega_V = (\tau_{sv}/\tau_{sv})^{-\sigma} \) and thus \( U \) is identical across all industries. If there is some industry for which \( \omega_V \) is close to one then reductions in TPU have a negligible direct effect on exporters in \( V \) and a negative indirect effect, as they face more competition from entry in other industries. Thus extending Proposition 2 as written requires \( \omega_V \) to be sufficiently far from unity for each \( V \) so we can order economic conditions in each \( V \) according to the policy state. Alternatively, we can extend the proposition to allow for the possibility just described.

\(^{38}\)We use the relationship \( c_V^U = \phi c_V^U \) and allow industry variation in the upgrade technology and sunk costs.

\(^{39}\)We show that proposition 1, 2 and 3 can be extended to this multi-industry setting and yield similar cutoff expressions, which now depend on industry specific trade costs.
3 Evidence

We use the model to examine how China’s WTO accession, which eliminated the annual MFN renewal debate in the U.S., contributed to its export boom to the U.S. We focus on industry exports, which will reflect both entry and upgrading effects, to estimate key structural parameters that allow us to perform counterfactuals. In section 3.8 we also quantify the impact of uncertainty on Chinese export entry and provide evidence for this channel in the Online Appendix C.10.

3.1 Approach

We estimate the export equation in changes for two reasons. First, it allows us to difference out unobserved industry characteristics such as entry costs, the productivity and mass of Chinese producers in \( V \) and the technology parameters in \( \zeta \). Second, we are interested in the impact of the change in uncertainty after the U.S. removed the threat of column 2 tariffs due to China’s WTO entry. So our baseline estimates employ a simple difference of \( \Delta \ln R = f(\bar{\omega}_V, \gamma) + b_\tau \Delta \ln \tau_V + b_d \Delta \ln D_V + b + e_V \) (24), where \( \Delta \ln x_V \equiv \ln x_{0V} - \ln x_{1V} \). The coefficient \( b_\tau = -\frac{k\sigma}{\sigma-1} < 0 \) is the effect of applied tariffs (conditional on the uncertainty factor). We model advalorem export costs, \( d_V \), as a function of observable shocks given by the advalorem equivalent of insurance and freight, \( \Delta \ln D_V \), an unobservable industry specific component that is differenced out, and an I.I.D error term contained in \( e_V \). The changes in transport cost allow us to identify the Pareto shape parameter, \( b_d = -k \). Any changes in aggregate expenditure on differentiated goods or its price index are captured in the constant term, \( b \).

The uncertainty effect is captured by \( f(\bar{\omega}_V, \gamma) = (k - \sigma + 1) \Delta \ln U_s(\frac{\tau_{2V}}{\tau_{1V}} g, \gamma) \). The model predicts that \( U_s \) changes over time if either tariffs or transition probabilities change. In the period we consider \( \tau_{2V}/\tau_{1V} \) is nearly constant within industries, so that \( f \) captures any changes in the probability of transition, \( \gamma \). The null hypothesis under the model is that prior to WTO accession there is a positive probability of column 2 tariffs, i.e. \( \lambda_{12} = \gamma \lambda_2 > 0 \), which disappears after accession. Therefore in the baseline \( f(\bar{\omega}_V, \gamma) = -(k - \sigma + 1) \ln U_1(\frac{\tau_{2V}}{\tau_{1V}} g, \gamma) \) and \( f \) is increasing in \( \tau_{2V}/\tau_{1V} \) if the WTO reduced \( \gamma \).

Standard trade models with a gravity structure yield a restricted version of (24) with \( f = 0 \) that is nested in our model. Even if uncertainty is important, our functional form assumptions may not be satisfied by the data. We address this as follows. First, we provide non-parametric estimates of the impact of \( \tau_{2V}/\tau_{1V} \) on export growth. Second, we control for observed changes in policies and trade costs and provide semi-parametric estimates of the impact of policy uncertainty—imposing only the gravity structure that is common in trade models without uncertainty. Third, we test the model’s functional form for \( f \) and perform numerous robustness checks (e.g. the possibility of industry-specific growth trends, unobserved demand and supply shocks, and other potential threats to identification). Fourth, we provide a non-linear structural estimate of \( f \) that we use to quantify the impact of TPU.
3.2 Data and Policy Background

We combine trade and policy data from several sources. Trade data for several years at the HS-6 digit level and U.S. tariffs are obtained via the World Bank’s WITS. The source of other policy measures we use are described in the Appendix [3]. The cost of insurance and freight from the U.S. Census is obtained via NBER. Tariff and transport cost data are converted to their iceberg form (e.g. from 10% to $ln\tau = ln1.1$).

There are 5,113 HS-6 industries in the 1996 classification; China exported in 3,617 of these in both 2000 and 2005 to the U.S. The baseline analysis focuses on the industries traded in both years so that a log growth rate exists. These industries account for 99.8% of all export growth from China to the U.S. in this period.

China’s accession to the WTO did not significantly change the applied trade policy barriers it faced in the U.S. [41] Thus our empirical analysis controls for changes in policy and export costs and focuses on the potential role of accession in reducing the policy uncertainty Chinese exporters faced in the U.S. While China was first granted MFN status by the U.S. in the 1980s, it was subject to annual renewal with severe consequences of revocation. China would have faced column 2 tariffs and a trade war would likely have ensued. For our baseline sample, which is summarized in Table 1, we calculate that the (simple) average tariff China would face in the U.S. would rise from 4% (MFN) to 31% (column 2) if it lost its MFN status in 2000. Although China never lost MFN status, it came quite close: in the 1990s Congress voted every year on whether to revoke MFN and the House passed such a bill three times.

The following policy background is useful in understanding the pre and post period choice for our baseline estimates. There was uncertainty about both China’s accession to the WTO and its permanent normal trade relations (PNTR) with the U.S. as late as 2000. [42] Protracted negotiations over China’s accession meant Congress voted again in the summer of 2001 over whether to revoke MFN. The president was required to determine whether the terms of China’s WTO accession satisfied its obligations under the Act. Otherwise the U.S. could opt-out of providing MFN status to China under Article XIII of the WTO, a right it had exercised with respect to other accessions. China joined the WTO on December 11, 2001 and the U.S. effectively enacted PNTR on January 1, 2002. [43] This strongly suggests that uncertainty about column 2 tariffs remained at least until 2000 and that it was not reduced until 2002. We will focus on the growth between 2000-2005 but show that the basic effect is present for other relevant periods.

3.3 Non-parametric and Semi-parametric Evidence

Table 1 provides summary statistics for our baseline sample. Export growth from 2000 to 2005 averaged 129 log points (lp) across HS-6 industries, with substantial variation across them: the standard deviation is 167 lp. This industry variation suggests that the boom can’t simply be explained by aggregate shocks. Table 1 also shows substantial variation in column 2 tariffs across the industries.

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40 Our baseline sample is smaller because it focuses on HS-6 with ad valorem tariffs. Tariffs in about 94% of HS-6 tariff lines in 2005, are levied on an ad valorem basis but some are specific tariffs levied on a per unit basis. We address this issue and the zero trade flow industries in the robustness checks.

41 The exception is textiles quotas that were fully lifted in 2005 and that we can control for empirically.

42 Foreign and economic relations between these countries remained tense into the late 1990s for several reasons. The Chinese embassy in Serbia was accidentally bombed by NATO in May 1999. Then in the summer of 2000 there was a vote in Congress to revoke China’s MFN status. In October 2000 Congress passed the U.S.-China Relations Act granting PNTR but its enactment was contingent on China’s accession to the WTO. In the meantime, a U.S. spy plane collided with a Chinese fighter jet over the South China Sea in April 2001.

43 Pregelj (2001) provides details on the U.S. MFN status relative to China.
All else equal, the model predicts lower initial export levels in the pre-WTO period for industries with higher potential profit losses if there was a possibility of tariff increases. If WTO accession reduces or eliminates this probability, we should observe relatively higher export growth in those industries. For any given value of $\sigma$ the industry ranking of potential profit loss is determined by $\frac{\tau_2 V}{\tau_1 V}$ so we use this ratio to partition the sample into the columns in Table 1 labeled low (bottom tercile of $\frac{\tau_2 V}{\tau_1 V}$) and high “uncertainty” industries. Export growth in high uncertainty industries is 18 log points higher, a mean difference that is statistically significant. In fact, the whole distribution of export growth for high uncertainty industries first order stochastically dominates the low uncertainty distribution according to a Kolmogorov-Smirnov test. These findings are robust to trimming the top and bottom percentile of each subsample.

Figure 4 provides further non-parametric evidence of this relationship by estimating a local linear regression (lowess) of export growth on $\ln \left( \frac{\tau_2 V}{\tau_1 V} \right)$. We confirm the higher growth in high initial uncertainty industries, as obtained in the mean test, and a non-negative relationship over the full domain.

Using a semi-parametric approach we can control for other potential determinants of export growth and also test for specific functional forms of the uncertainty term. Several trade models yield a gravity equation that is a special case of (24) with the implicit restriction that $f = 0$. We use the residuals from that restricted estimation to determine how $\frac{\tau_2 V}{\tau_1 V}$ affects $f$ without imposing functional forms on the latter. Using a double residual semi-parametric regression (Robinson, 1988) we find that $\frac{\tau_2 V}{\tau_1 V}$ has a significant effect on subsequent export growth net of tariff or transport cost changes. This result is robust to including 21 sector dummies in the restricted regression to net out any heterogenous growth trends in similar HS-6 industries.

In Figure 5, we plot the resulting semi-parametric fit (that did not impose any $\sigma$) and see it is increasing in $1 - \left( \frac{\tau_2 V}{\tau_1 V} \right)^{-3}$—the potential profit loss measure without GE effects when $\sigma = 3$. The predicted parametric line obtained from OLS estimation of (24) when the loss measure is $1 - \left( \frac{\tau_2 V}{\tau_1 V} \right)^{-3}$ lies everywhere within the semi-parametric 95% confidence interval. A formal test reveals that we fail to reject the equality of this particular parametric fit and the semi-parametric one. Since we can’t reject a linear fit with the parametric loss measure, we use OLS specifications as a baseline to perform various robustness tests. Choosing, $\sigma = 3$ is consistent with the median value for the U.S. estimated by Broda and Weinstein (2006) and robust to alternative choices of the elasticity of substitution. We also test if the semi-parametric fit is equal to alternative parametric fits that are linear or log linear in $\tau_2 V / \tau_1 V$ and find they are rejected in the data. This suggests that reduced form measures of column 2 tariffs should not be used for quantitative predictions. In part, this is because the non-linearity implies that the marginal effect of $\tau_2 V$ is smaller at high tariffs that could be well above trade prohibitive levels.

### 3.4 Linear Estimates

The semi-parametric evidence supports approximating the uncertainty term using $b_\gamma \times \left( 1 - \left( \frac{\tau_2 V}{\tau_1 V} \right)^{-3} \right)$ in (24). When we approximate $U_1 (\tilde{\omega}, \gamma)$ around $\gamma = 0$ and use (15) we have the following structural interpretation of $b_\gamma^{OLS} = \frac{k-\sigma+1}{\sigma} u(\gamma) g \geq 0$. So we first present parametric estimates of $b_\gamma$ and examine their robustness to two potentially important sources of omitted variable bias.

**Baseline:** The OLS results in Table 2 are consistent with the structural interpretation of the parameters. In column 1 we see that $b_\gamma$ is positive and significant. As predicted, this implies that industries with higher initial potential losses grew faster after WTO accession. The coefficients on tariffs and transport costs are
negative and significant. The estimation equation contains an over identifying restriction, \( b_\tau = \frac{\sigma}{\sigma - 1} b_d \), that we can’t reject. We therefore re-estimate the model in column 2 with this restriction, which increases the precision of the estimates.

**Sector level growth trends and unobserved heterogeneity:** The model contains several unobservables that can vary across industries. Recall that most of these are time invariant and log separable (e.g. sunk costs, upgrade technology, etc and the mass of Chinese firms) and thus are already differenced out in the baseline estimates. Any growth innovations common to all industries are absorbed in the baseline constant, \( b \). We now allow for that growth to differ across sectors by including a full set of 21 sector dummies in the differenced equation \(24\). The results shown in columns 3 and 4 of Table 2 are similar to those in the baseline—importantly, \( b_\tau \) remains positive and significant. This specification controls for several potential sources of omitted variable bias, such as differential changes across sectors in productivity, sunk costs, FDI and Chinese barriers on intermediates\(25\).

**Non-tariff barriers:** The regressions in columns 3 and 4 of Table 2 also control for any sector level changes in non-tariff barriers (NTBs). Nevertheless, some of those barriers can also vary at the industry (HS-6) level. We address this with binary indicators for whether an industry had any of the following barriers in a given year: anti-dumping duties, countervailing duties and China-specific special safeguards. Following China’s accession to the WTO it also became eligible to benefit from the phase-out of quotas in textiles that had been agreed by WTO members prior to China’s accession under the Multi-Fiber Agreement (MFA), which was fully implemented by 2005. We have indicators for the HS-6 industries where such quotas were lifted\(25\).

In **Table 3**, we compare the baseline results (column 1) with alternative specifications that control for changes in NTBs and in MFA quotas. In column 2 we include a regressor for the change in the binary indicator for both MFA quotas and NTBs and find they have the expected negative sign\(26\). Their inclusion does not affect the other coefficients whether or not we control for sector effects.

NTBs may respond to import surges from China. To the extent that these surges are more likely in some sectors, our sector effects in column 3 already control for this potential endogeneity\(27\). To address the possibility that this reverse causality could also occur within sectors, we instrument the change in NTB with its level binary indicator in early years—1997 and 1998. Column 4 shows that instrumenting does not affect the coefficient for uncertainty relative to the OLS version (column 3 of Table 3) or the specification without the NTB variable (column 3 of Table 2)\(25\).

### 3.5 Robustness, Falsification Tests and Additional Evidence

**Robustness**

\(44\)The coefficients on tariffs and transport costs are significant in the constrained regressions. One reason for the increase in precision is that most applied tariff changes are very small during our sample period and there may be a few influential observations. We address this with robust regression methods in Table A2 and find results that are qualitatively similar to Table 2 with statistically significant estimates for the uncertainty and tariff coefficients.

\(45\)Additional details on the TTB and MFA indicators and sources appear in the data Appendix B.

\(46\)Below we also provide panel evidence that the baseline results in 2000-2005 are similar to those in 2000-2004, which was a period when the quotas were mostly still in place. The panel results are robust to dropping products that ever had an MFA quota regardless of the year it was removed.

\(47\)The MFA dates back to the 1980s and its phaseout was implemented with the Uruguay Round in 1996 before China was a member of the GATT/WTO. As such, it is plausibly exogenous as a barrier to China’s imports.

\(48\)The two instruments pass a Sargan over-identifying restriction test and we also fail to reject the exogeneity of the TTB variable using a Durbin-Wu-Hausman test. The instruments have significant explanatory power in the first stage, with the relevant F-statistic above 10. We also find that the constrained version \( (b_\tau = \frac{\sigma}{\sigma - 1} b_d) \) yields very similar coefficients for the uncertainty, tariff and transport variables if we include the NTB and MFA (column 5 of Table 3) or not (column 4 of Table 2).
Table A2 summarizes the robustness of the baseline linear estimates of $b_\gamma$ (replicated in the first two unnumbered columns for comparison). The specifications also include tariff and transport cost changes as well as a constant or sector effects, which are not reported due to space considerations. The central point is that the sign and significance of $b_\gamma$ in the baseline are robust to the following potential issues:

- **Alternative elasticity of substitution.** The semi-parametric evidence suggests $\sigma = 3$ is a reasonable value and this is also the median value for the U.S. estimated by Broda and Weinstein (2006). In columns 1-4 of panel A we use $\sigma = 2, 4$ to compute the uncertainty measure. To address the possibility that some industries have elasticities very different from the overall median we do the following. Let $\tilde{\sigma}_V$ denote the median HS-10 elasticity estimate (from Broda and Weinstein) in any given HS-6 industry. In column 5 we drop observations with $\tilde{\sigma}_V \notin [1.5, 4.5]$. In column 6 we use the terciles of $\tilde{\sigma}_V$ to distinguish between low, medium and high elasticity industries and use the median within each of these bins to recompute the uncertainty measure.

- **Potential outliers.** In columns 1 and 2 of Panel B we employ a robust regression procedure that downweights outliers.

- **Sample selection.** Over 98% of export growth occurred in the industries contained in the baseline sample. However, when there is no trade in 2000 or 2005 we can’t compute log growth, which reduces the sample. In columns 3 and 4 of Panel B we address this by using mid-point growth as our dependent variable. Another source of sample selection in the baseline is the exclusion of industries that only have specific tariffs. Columns 5 and 6 use these additional industries by calculating advalorem equivalents (given by specific tariff $V_{unit value}$) and incorporating them into both the change in applied tariffs and uncertainty.

- **Processing trade.** Chinese exports in certain industries reflect mostly processing trade – foreign firms supply inputs and parts that are assembled in China and returned to the foreign firm (Kee and Tang, 2013). If our results were driven by these industries then they could reflect other factors such as changes in Chinese policies towards processing trade. In columns 7 and 8 of Panel B we drop all the HS-6 industries in the section of the Harmonized System with the largest share of processing trade.

**Falsification Tests**

We addressed omitted variable bias thus far by controlling for specific variables at the HS-6 level and unobserved contemporaneous sector shocks. In Table A3 we provide evidence against bias from unobserved industry demand and supply shocks.

Suppose there was an unobserved shock to Chinese production (and/or consumption) that was correlated with our measure of TPU. In that case our baseline estimates would be biased. If this was a China specific shock then it would affect its exports to all markets, particularly those with similar size and income per capita as the U.S.. We test this in column 2 by regressing Chinese exports to the European Union on U.S. TPU and find a point estimate close to zero and insignificant.

49 The typical coefficient for transport cost is $b_d = -2.5$ and $b_\tau = b_d \sigma / (\sigma - 1)$ can’t be rejected at p-values listed in last row. All the results are available upon request.

50 Section XVI: Machinery and Mechanical Appliances; Electrical Equipment; Parts Thereof; Sound Recorders and Reproducers, Television Image and Sound Recorders and Reproducers, and Parts and Accessories of Such Articles

51 Column 1 shows this specification for the U.S. applied to the common subsample of industries exporting to both destinations to avoid any sample composition issues. The other difference relative to the baseline in column 3 of Table 2 is that we do not include the transport cost since we do not have that data for the E.U.
If U.S. production decreased (and/or its consumption increased) in industries where China faced higher initial uncertainty then the baseline estimates would tend to be biased up. Such shocks would also increase U.S. imports from other countries. In column 4 we see that is not the case: U.S. imports from Taiwan were not significantly different in industries with higher initial TPU.

The falsification test for Taiwan keeps a number of important factors constant. First, Taiwan faced a change in U.S. import policy very similar to China’s in this period: WTO accession. Second, prior to accession Taiwanese exporters faced MFN tariffs in the U.S. and if they had lost MFN status they would have faced the same column 2 threat tariffs as China’s exporters. But Taiwan was never subject to an annual renewal process for its MFN status so the model would predict little or no change in the probability of losing MFN status upon accession, and the results in column 4 support this prediction.

Robustness to pre-accession industry growth trends

Pre-accession growth trends could also generate an omitted variable bias, if they continued post-accession and were correlated with the uncertainty measure. We examine this possibility by first running our baseline estimation on pre-accession Chinese import growth. In Table A4, column 3 we find no significant effect of the uncertainty measure in 1996 on export growth in 1996-99. To eliminate any HS6 industry growth trends that persist from the pre- to the post-accession period we subtract the pre-accession equation (in changes) from the baseline post-accession equation (also in changes). This difference of differences identification approach is similar to Trefler (2004) so we relegate the econometric details to Online Appendix C.8. Columns 1 and 2 of Table A4 show the baseline results are not driven by pre-accession growth trends.

Yearly panel evidence and timing of TPU change

The results thus far focus on specific years and a balanced panel. In our working paper we explore the full panel to examine if the uncertainty coefficient changed over time. For comparison to our earlier results we hold fixed the profit loss measure calculated using applied tariffs from 2000. In Online Appendix C.9 we show that the structural interpretation of the full panel estimates is \( \hat{\gamma}_{\text{panel}} = \frac{k-\sigma+1}{\sigma-1} \frac{\beta \lambda_2}{1-\beta \lambda_2} \Delta \gamma_t \), where \( \Delta \gamma_t = \gamma_{2000} - \gamma_t \) for any pre-accession year \( t \). The estimates are plotted in Appendix Figure A1 and show no significant change in TPU in the pre-accession period: 1996-2001. This indicates that minor changes in the legislation or in the relations between the U.S. and China did not significantly affect Chinese firms’ beliefs about losing the MFN status. Those beliefs seem to have been revised only after China accedes to the WTO and obtains PNTR. From 2002-2005 we find a positive and significant coefficient and its magnitude in 2005 is similar to the baseline. This timing evidence indicates that accession did lower uncertainty as predicted by the model.

3.6 Non-linear Structural Estimates

To quantify the effect of policy uncertainty we focus on non-linear estimates of the model’s structural parameters. We do so because the linear approximation could and does overstate the impact of uncertainty since \( U \) is concave in \( \gamma \), as we show below.

If there is a negligible probability of column 2 tariffs after WTO accession, then \( f \left( \hat{U}_V \right) = -(k-\sigma+1) \ln U_1(\hat{\omega}_V, \gamma) \). Using the definition of \( U_1 \) to rewrite \( \hat{U}_V \) in terms of estimable coefficients \( b_x \) we obtain
\[
\Delta \ln R_V = b_{d\sigma} \ln \left[ 1 + \frac{\hat{b}_\gamma \left( \frac{\tau_{UV}}{\tau_V} \right)^{-b_\sigma}}{1 + \hat{b}_\gamma / g} \right] + b_r \Delta \ln \tau_V + b_d \Delta \ln D_V + b + \epsilon_V
\]

where \( b_d = -k \) and \( b_r = -\frac{k\sigma}{\sigma - 1} \), as before. We further define \( \hat{b}_\gamma = u(\gamma) g \), \( b_\sigma = \sigma \), and \( b_{d\sigma} = -\frac{k - \sigma + 1}{\sigma - 1} \). One component of \( U_1 \), \( 1 + \hat{b}_\gamma / g \), is log-separable and does not vary by industry. We cannot identify it separately from the constant term, \( b \), but we will be able to compute it from our estimates for the quantification.

The non-linear baseline imposes two of the theoretical restrictions: \( b_{d\sigma} = \frac{(b_d + b_\sigma - 1)}{b_\sigma - 1} \) and \( b_r = \frac{b_d}{b_\sigma - 1} \), and estimates \( \hat{b}_\gamma \), \( b_d \) and \( b \) (we initially restrict \( b_\sigma = 3 \) and then test it). Column 1 of Table 4 provides non-linear least squares (NLLS) estimates. For comparison with earlier results we transform the estimate for \( \hat{b}_\gamma = 0.687 \) into its OLS regression counterpart, \( b_\gamma^{\text{NLLS}} = \frac{b_d - b_\sigma + 1}{b_\sigma - 1} \hat{b}_\gamma = 0.82 \). This is slightly higher than its OLS counterpart (column 2) and significantly different from zero. That is also the case after we control for sector effects (comparing columns 3 and 4).

**Consistency with model and other evidence**

Before using these estimates for quantification we provide evidence that they are consistent within the model and relative to other evidence. The signs of all the estimated parameters are those predicted by the model. We re-ran the baseline specifications by relaxing each of the two restrictions individually, either \( b_r = b_d \frac{\sigma}{\sigma - 1} \) or \( b_\sigma = 3 \). We fail to reject these restrictions and report p-values in the last two rows of Table 4. Our choice of \( \sigma = 3 \) is typical in trade estimates and the partial elasticity of exports to tariffs in the absence of uncertainty, -6.6, is close to previous ones that use similarly disaggregated U.S. trade and tariff data.

Under a Pareto productivity distribution with shape parameter \( k \), export sales without uncertainty are also Pareto but with shape \( k / (\sigma - 1) \). The 95% confidence interval for our estimate of the sales distribution parameter is \( b_d / (\sigma - 1) \) is [1.4, 3.0]. The estimate is larger than 1 and satisfies the model requirement for a finite first moment of sales that we did not impose. The magnitude is similar to what is found by other studies using firm level data.

The other parameter central to the quantification is \( u \)—a new Chinese exporter’s expected duration of a spell under column 2 prior to WTO accession. We can back out an estimate of \( u \) from our estimates. The expected duration is a belief held by the exporters for an event that never took place so there is no way to defend a particular value. Nevertheless, the bounds on our estimate are reasonable and consistent with the model. We obtain an upper bound for the average duration \( \bar{u} = \hat{b}_\gamma / \bar{g} = .64 \) where \( \bar{g} \) is the upper bound for the ratio of the MFN and worst case scenario price index in \( \Delta \). In Appendix C.1.1 we show how to use the NLLS estimates and data on Chinese import penetration to obtain \( \bar{g} = (P^D \circ P^D)^{-1} \approx 1.038^2 \).

To interpret the duration estimate consider a Chinese firm that starts exporting in 2000. Its expected export survival would be 6.25 years using an exit rate of 0.16 given by the fraction of new Chinese exporters that stop exporting after one year reported in Ma et al. (2014). These values imply a new exporter expected

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52 Given that the NLLS estimation relies on the model structure and the variation in the transport cost variable to identify \( k \), we minimize the potential influence of outliers by focusing on the subsample without transport cost outliers, as measured by changes in costs more than three times the interquartile range value beyond the top or bottom quartile value of the baseline sample. The estimate for \( k \) in this subsample is higher (under NLLS or OLS) than the baseline, which suggests that the transport cost for some products contained measurement error and generated attenuation bias.

53 Romalis (2007) estimates this elasticity to be between 6.3 and 6.7 using US statutory tariffs and HS-6 imports.

54 Eaton et al (2011) obtain an aggregate estimate of 2.46 using French exports; Hsieh and Ossa (2011) obtain a range from 1 to 1.44 over industries using Chinese firm data.
to spend at least 10% of their exporting spell under column 2. The upper bound in 2000 is 40% but only under the implausible belief of a zero probability of entering the WTO. Furthermore, from the definition of $\hat{b}_\gamma$ and $u$ we obtain an estimate for the transition probability $\hat{\lambda}_{12} = \frac{1 - \beta \hat{\lambda}_{22} \hat{b}_\gamma}{\beta g}$. While the estimation does not impose any restriction on this parameter, we find that it is bounded in the unit interval for any $\beta > .687$, i.e. for any reasonable rates of firm annual survival.\footnote{55}

3.7 Quantification and Counterfactuals

We use the NLLS estimates in Table 4 to quantify the effect of policy uncertainty changes and decompose it into a pure risk and mean effect.

Table 8 summarizes the key parameters used in each exercise. We stress three points. First, we do not require or use any assumption on the values of the following individual parameters $\beta$, $\lambda_{22}$ or $\lambda_{12}$ since all the relevant information is contained in $u$. Second, we will not calibrate any parameters, we only use parameters that we estimated (or parametrized and tested in the case of $\sigma$) and combine them with any data required by the model for aggregating across industries. Third, price index effects require a measure of China’s share in U.S. expenditure on differentiated goods. In the model that expenditure is endogenously given by fixed nominal worker income. To relax this assumption we use data on U.S. manufacturing expenditure to compute counterfactuals that involve price effects.\footnote{56}

Partial and General Equilibrium Effects of Uncertainty on Exports

Given our estimates, what would be the average impact of re-introducing TPU in a post accession year $t$ on any given outcome in that year (trade, prices, welfare)? In Table 5 we compute the difference between the predicted effect without uncertainty and this counterfactual and label it the contribution of the policy uncertainty reduction.\footnote{57}

The following average over industries of the first term in (25) captures one central part of the uncertainty effect on export growth:

$$ E_V \left( \ln R_{0V} - \ln R_{1V} \right) |_{r, \sigma, \beta} = b_{d\sigma} E_V \left[ \ln \frac{1 + \hat{b}_\gamma \left( \frac{R_{2V}}{R_{1V}} \right)^{-\sigma}}{1 + \hat{b}_\gamma / g} \right] $$

(26)

The contribution of TPU if there was no GE effect is obtained by evaluating the expression at $g = 1$. In column 1 of Table 5 we use the NLLS estimates and find that if the pre-WTO uncertainty was re-introduced in 2005, China’s exports would fall by 29 log points.

Next, we capture price index effects through two components. First, we evaluate (26) and incorporate the price effects in $g$ by using the upper bound $\hat{g} = 1.08$. We obtain 26 log points of export growth due to what we label the Partial GE effect of TPU in column 2, Table 5. Second, in column 3 we augment the

\footnote{55}If the probability of entering into the WTO is zero then there are only two relevant policy states and starting at MFN the expected time under column 2 is simply $u / (1 + u) \leq 0.4$ for all $g \geq 1$.

\footnote{56}To see this note that $\hat{\lambda}_{12} \in (\frac{1 - \beta g \hat{b}_\gamma}{\beta g}, \frac{1 - \beta \hat{b}_\gamma}{\beta g})$ where the lowest value occurs if column 2 tariffs are an absorbing policy state, $\lambda_{22} = 1$, and the highest if $\lambda_{22} = 0$ and $\hat{g} = 1$.

\footnote{57}Over 98% of Chinese exports to the US are in manufactures.

\footnote{58}Since we are trying to capture the effect of a reduction in uncertainty relative to outcomes where uncertainty remains high, we ignore the transition effects associated with increases in uncertainty.
partial effect with the contemporaneous price impact of changing uncertainty:

\[
E_V (\ln R_{0V} - \ln R_{1V}) |_{\bar{\tau}, \bar{D}} = E_V (\ln R_{0V} - \ln R_{1V}) |_{\bar{\tau}, \bar{D}, \bar{R}} + k (\ln P_0/P_1) |_{\bar{\tau}, \bar{D}}
\]

Using the model structure, data and estimated parameters we calculate the growth in the price index due to a change in uncertainty starting from a post-accession period without uncertainty: \(\ln \left[ P_1/P_0 \right]_{\bar{\tau}, \bar{D}} \approx \gamma d \ln P_1 |_{\gamma=0}\) (see Appendix C.1.1). This price change is about 0.9 log points so \(k (\ln P_0/P_1) |_{\bar{\tau}, \bar{D}}\) is -4.0 log points. Adding this to the effect in column 2 in Table 5 we obtain the **Full GE effect** of 22 log points.

Accounting for the non-linear model structure and GE effects is quantitatively important. The change in imports implied by the OLS estimates in column 2, Table 4 is \(b_{OLS} \gamma \times E \left( 1 - \left( \frac{\tau_2 V}{\tau_1 V} \right)^{-3} \right) = 34\) log points. Since aggregate imports from China in 2005 in our sample were about $250 billion, the OLS estimates without GE factors predict an effect of TPU of $72 billion in 2005, whereas it is $50 billion under the NLLS estimates with GE effects.

Using the NLLS structural parameters we can also compute counterfactual import values for other years. These are presented in **Figure 1** which normalizes imports by U.S. expenditure on manufacturing to provide the Chinese import penetration ratio. The solid line shows this ratio tripled between 2000 and 2010. The dashed line is the full GE counterfactual if uncertainty in any year from 2002-2010 returned to its 2000 level. Under TPU this ratio would have been 3.6% in 2005 instead of 4.5% and in 2010 it would have been 5.5% instead of 6.7%.

**Consumer welfare, profits and employment trade cost equivalents**

We now quantify the effect of TPU on the welfare of consumers identified in proposition 3. The welfare change in the current state is \(-\mu \gamma d \ln P |_{\gamma=0}\), which is about 0.8 percent.\(^{59}\) The magnitude of this welfare effect of TPU on trade is comparable to estimates arising from alternative channels. For example, Broda and Weinstein (2006) estimate that the real income gain from new imported varieties in the U.S. between 1990-2001 was 0.8 percent.\(^{60}\) Costinot and Rodríguez-Clare (2014) calculate that a worldwide tariff war (uniform tariffs of 40%) would lower North American welfare by 0.7 percent in a static model with heterogenous firms, monopolistic competition and multiple sectors.\(^{61}\)

We can also quantify the importance of TPU via a trade cost equivalent. This amounts to determining what change in average trade costs under certainty is necessary to generate the same impact as eliminating TPU. We compute this by dividing the effect of uncertainty on trade from column 3 of Table 5, 22 log points, by the deterministic trade tariff elasticity, including the partial effect, \(-k \sigma \sigma^{-1} = -6.6\), and the effect via the price index, \(k x \frac{d \ln P}{d \ln \tau V} = 4.4 \times 0.12\), which we derive in the Online Appendix C.1.2. Column 1 of Table 6 shows that in 2005 the export impact of increasing TPU to its 2000 level was equivalent to increasing tariffs by 3.6 percentage points. Column 2 shows the tariff equivalent of TPU is twice as high for Chinese export entry, which we will discuss further below.

---

\(^{59}\)We uses \(\mu = 0.86\) and estimate a price effect of about 1 percent (see Table 8 and Online Appendix C.1.2).

\(^{60}\)As is standard in most trade models neither of these quantifications takes into account services. However, the model and calculations do take into account the large fraction of non-traded goods since many of the differentiated goods are produced by firms that are not productive enough to export. This is reflected in the low values of import penetration from China that reflect U.S. imports/consumption.

\(^{61}\)Our model differs on some important dimensions: e.g. uncertainty, sunk costs, an outside good and no free entry in the domestic market.
Column 3 of Table 6 presents trade cost equivalents of TPU for the price index of 7.6 percentage points for the tariff, which are obtained using

\[ \Delta_\tau = \gamma \left[ \frac{d \ln P}{d \gamma} \right] \left[ \frac{d \ln P}{d \ln \tau_V} \right]_{\gamma=0} \]  

This value also applies to domestic profits, sales and employment of U.S. firms in the differentiated sector since these outcomes depend on both uncertainty and import tariffs only via \( \tau \). This is also the case for consumer welfare in the MFN state so the impact of re-introducing TPU in 2005 is also equivalent to increasing tariffs by 7.6 percentage points. The second row of Table 6 provides similar calculations for transport costs that are higher than for tariffs because the latter have a stronger effect on trade.

Re-introducing TPU in 2005 would not only reduce welfare at given tariffs but also change expected future welfare if the tariffs eventually changed. Since the MFN and agreement tariffs were not very different, the additional effect mainly reflects the welfare cost of a spell of duration \( u \) under column 2 tariffs when the price index is about 3.8 percent higher. Thus in addition to the 0.8 percent welfare reduction in the current state, TPU would lower expected welfare by as much as 2 percent.\(^{63}\)

**Mean-risk decomposition**

We defined TPU as the probability of a tariff change starting at the MFN state, \( \tau_1 \), and parametrized it by \( \gamma \). We can decompose the uncertainty effect into a risk and a lock-in component as follows

\[ \ln \frac{R_V(\gamma_0, \tau_{1V})}{R_V(\gamma_1, \tau_{1V})} = \ln \frac{R_V(\gamma_0, \bar{\tau}_V)}{R_V(\gamma_0, \bar{\tau}_V)} + \left[ \ln \frac{R_V(\gamma_0, \tau_{1V})}{R_V(\gamma_0, \bar{\tau}_V)} - \ln \frac{R_V(\gamma_1, \tau_{1V})}{R_V(\gamma_1, \bar{\tau}_V)} \right] \quad \text{each } V \]  

where \( \bar{\tau}_V \) is the long-run mean of the tariff in each industry. The first term is the growth in exports due to credibly securing tariffs at their long-run mean, i.e. a risk reduction due to a mean preserving compression in tariffs. The second term in brackets is positive when the initial tariffs are below the long-run mean and the agreement locks in lower tariffs, as is the case in our application.

We quantify the risk reduction term, by using the full GE effect expression in (27) evaluated at the long-run mean of the tariff starting at MFN to be independent of \( \gamma \) then we need to assume that both the WTO state and column 2 states are absorbing.\(^{64}\) Second, we must consider alternative odds of the WTO relative to column 2 that are independent of the value of \( \gamma \); if we assume \( \lambda_{10}/\lambda_{12} = 2 \) then \( \bar{\tau}_V \) ranges from 1 to 2.04 across industries and its overall average is 1.15.

Using these counterfactual mean tariff levels and our parameter estimates we compute counterfactual levels of the price index, trade and import penetration to calculate \( \ln \frac{R_V(\gamma_0, \bar{\tau})}{R_V(\gamma_1, \bar{\tau})} \). The results in Table 7 show that the average export growth due to the risk reduction is 82% of the 22 log points for the full GE effect. This risk share is large because the threat of moving to \( \tau_2 \) entails a doubling of tariffs even if we start at the mean. For alternative odds of entering the WTO vs. column 2 of 1:1 and 3:1 we find risk shares between 65

\(^{62}\)We can also use (20) and the parameter estimates to compute the expected welfare effect when \( \tau_1 \approx \tau_0 \) and the implied \( \Delta_\tau \) for it is 8 percentage points.

\(^{63}\)We derive the general expression in Appendix C.4 and show \( u \times \mu \times \ln P^*_1/P^*_0 = 0.64 \times 0.86 \times 0.037 \).

\(^{64}\)In this case if we start at \( \tau_{1V} \) the long-run mean is simply \( \bar{\tau}_V = \frac{\lambda_{10}}{\lambda_{10}+\lambda_{12}} \tau_{1V} + \frac{\lambda_{12}}{\lambda_{10}+\lambda_{12}} \tau_{2V} \), reflecting the probability of going into an agreement relative to a trade war conditional on abandoning the temporary MFN policy.

\(^{65}\)See Online Appendix C.4 for more details on this procedure.
and 88 percent of the total effect. A similar exercise for the price index and the welfare effect shows that the risk reduction component is about 43% of the total. This is smaller than the export effect because the price index effect depends critically on import penetration, which is considerably lower at the counterfactual mean tariff (2.6%) than in the data (4.5%).

### 3.8 Entry

We now examine the effect of uncertainty on Chinese export entry. We first explore the structural estimates to quantify the role of entry. In Online Appendix C.10 we corroborate this evidence using detailed information on new varieties exported from China to the U.S. (Table A5).

The model predicts that out of all Chinese firms in industry \( V \) at least a fraction \( G(V) \) export to the U.S. By employing the same cutoff expression and Pareto distribution used in the export equation we obtain the following expression for Chinese net entry into exports to the U.S., between 2005 and 2000:

\[
\Delta \ln n_V = -k \ln U_{1V} - \frac{k\sigma}{\sigma - 1} \Delta \ln \tau_V - k\Delta \ln D_V + k\Delta \ln P + b + \varepsilon_V \tag{30}
\]

where \( b = \frac{k}{\sigma - 1} \Delta \ln E \). Note that the elasticity of entry with respect to \( U \) is higher than in the export equation by a factor \( k/(k - \sigma + 1) \) but all other elasticities are unchanged. We use these relationships between coefficients to quantify the uncertainty impact on entry in Table 5. For the partial effect for example, we calculate

\[
\mathbb{E}_V (\ln n_{0V} - \ln n_{1V}) \big| \bar{\tau}, \bar{D}, \bar{P} = \frac{k}{k - \sigma + 1} \mathbb{E}_V (\ln R_{0V} - \ln R_{1V}) \big| \bar{\tau}, \bar{D}, \bar{P}
\]

As before, we focus on the NLLS estimates (column 1 Table 4) and calculate the contribution of uncertainty reduction to Chinese variety growth as the difference between the predicted effect without TPU and the counterfactual average impact of re-introducing TPU in 2005.

The full GE effect of TPU on Chinese variety growth is 43 log points, which is sizeable relative to the total growth in the number of Chinese exporting firms to the world over 2000-2005 (83 log points, Ma et al., 2014). In Table 7 we show that most of this effect (34 log points) is attributable to a risk reduction. In Table 6 we provide the trade cost equivalents for exports. Because the effect of TPU on entry is roughly twice that on export values so are the trade cost equivalents of uncertainty on entry: 11% for transport and 7.1% for tariffs.

The results are similar if we use a geometric mean of the tariffs to calculate the long-run mean.

The welfare effect we re-compute \(-\mu \gamma \frac{1}{\gamma} \ln \bar{P}_{\gamma=0} \) evaluated at the mean tariff. This also provides an upper bound for the first order risk effect of TPU on expected welfare. We show in the online Appendix C.4 that this occurs because the second term in (20) is non-negative to a first order approximation when evaluated at \( \bar{\tau} \) since per period welfare in the deterministic states is approximately log linear in tariffs in the relevant range.

If at \( s \) the current economic conditions are no worse than in the past, then \( n_{sV} \) is exactly equal to that fraction.

66 The results are similar if we use a geometric mean of the tariffs to calculate the long-run mean.

67 For the welfare effect we re-compute \(-\mu \gamma \frac{1}{\gamma} \ln \bar{P}_{\gamma=0} \) evaluated at the mean tariff. This also provides an upper bound for the first order risk effect of TPU on expected welfare. We show in the online Appendix C.4 that this occurs because the second term in (20) is non-negative to a first order approximation when evaluated at \( \bar{\tau} \) since per period welfare in the deterministic states is approximately log linear in tariffs in the relevant range.

68 If at \( s \) the current economic conditions are no worse than in the past, then \( n_{sV} \) is exactly equal to that fraction.
4 Conclusion

We assess the impact of trade policy uncertainty in a tractable general equilibrium framework with heterogeneous firms. We show that increased policy uncertainty reduces investment in export entry and technology upgrading, which in turn reduces trade flows and real income for consumers. We apply the model to the period surrounding China’s accession to the WTO. China’s WTO membership lead the U.S. to grant it permanent most-favored-nation tariff treatment, ending the annual threat to revoke MFN and subject Chinese imports to Smoot-Hawley tariffs. While recent work focuses on the costs of the Chinese export boom to employment and wages, we focus on the potential for gains from reducing TPU.

We derive observable, theory-consistent measures of TPU and estimate its effect on trade flows, prices and welfare. Had MFN status been revoked, the typical Chinese exporter would have faced an average tariff of 31%. According to our most conservative estimates this threat had large effects on trade and if it was re-imposed in 2005 it would have lowered Chinese exports by at least 22%. The welfare cost of this uncertainty was at least 0.8 percent of consumer real income, or the equivalent of a permanent tariff increase of 8 percentage points on Chinese goods. This is a substantial amount of effective protection, especially relative to the average applied tariffs, which were only about 4 percent.

Our findings have implications beyond this particular important event. They also indicate that an important role of agreements is to reduce policy uncertainty. More generally, our research points to substantial effects of policy uncertainty on economic activity and to the value of specific data-rich trade settings to identify them.

References


A Theory Appendix

A.1 Proof of Proposition 1

To prove the uniqueness of the industry cutoff in part (1a), we first establish sufficient conditions for a unique tariff below which each firm enters.

Lemma 1 (Entry threshold): For any given policy regime $\Lambda (\tau_m, \gamma)$ that exhibits uncertainty persistence, and each firm $c$ from a small exporting country, there is a unique threshold tariff per state, $\tau_{U}^{c} (\gamma, c)$, below which a firm enters into exporting.

Proof: Rewriting (2) recursively we have

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

Substitute in (4) to obtain

$$\Pi(a_s, c, \gamma) - \Pi_e(a_s, c, \gamma) + K = max \{0, \beta \mathbb{E}_s [\Pi(a_s', c, \gamma) - \Pi_e(a_s', c, \gamma)] - \pi(a_s, c) + K\}$$

where the option value of waiting is $V_s \equiv \Pi(a_s, c, \gamma) - \Pi_e(a_s, c, \gamma) + K$ and $\mathbb{E}_s V_s' \equiv \mathbb{E}_s [\Pi(a_s', c, \gamma) - \Pi_e(a_s', c, \gamma) + K]$.

(1) Entry by firms from small exporting countries have no effect on the importer aggregates. Thus for given $E$ and $P$ we have $a_s \equiv EP^{\sigma - 1} \tau_s - \sigma \sigma (\sigma - 1)^{\sigma - 1}$ so $\mathbb{E}_s V_s' = \int V_s d\Lambda (\gamma, \tau'|\tau)$.

(2) Because $-\pi(a_s, c)$ is increasing in $\tau$ it is more attractive to wait at higher tariffs because second element of (32) and therefore $V_s$ would be higher, all else equal.

(3) Since $\Lambda$ exhibits uncertainty persistence we have $\int V_s d\Lambda (\gamma, \tau' + \varepsilon) > \int V_s d\Lambda (\gamma, \tau'|\tau)$ if $V_s$ is increasing in $\tau$.

Given (3) if we start with an increasing $V_s$ the fixed point to this iteration is also increasing in $\tau$. By properties (2) and (3), $\beta \mathbb{E}_s V_s' - \pi(a_s, c)$ is increasing in $\tau$, so there is some $\tau_{U}^{c} (\gamma, c)$ below which the firm
value is higher if exporting and above which the opposite is true. QED

Proof of Prop. 1(a): $c^U_s$ is unique and $c^U_s = c^D_s U_1 (\omega, \gamma)$

Lemma 1 shows that each firm $v$ has a single tariff entry cutoff $\tau^U_s (\gamma, c_v)$. All firms have different cost but face the same $\tau$ and $\gamma$ in the industry so there is a unique entry cutoff for any given $\tau_m$, $c^U_s (\tau_m, \gamma)$, and only those with cost below this enter into exporting.

To show $c^U_s = c^D_s U_1 (\omega, \gamma)$ we first derive $\Pi(a_s, c \leq c^U_s, \gamma)$ if $s = 1$. Starting with (32) and taking the expectation over the possible states we have

$$\mathbb{E}_s V'_s = \lambda_{s+1} \left[ \beta \mathbb{E}_{s+1} V'_s - \pi (a_{s+1}, c) + K (1 - \beta) \right] \quad \text{if } c \leq c^U_s$$

$$= \lambda_{s+1} \left[ \frac{\lambda_{s+1} \pi (a_{s+1}, c) + K (1 - \beta)}{1 - \beta \lambda_{s+1} (1 + \beta \lambda_{s+1})} \right]$$

where the second line uses (32) and takes the conditional expectation starting at $s + 1$:

$$\mathbb{E}_{s+1} V'_s = \lambda_{s+1} \left[ \beta \mathbb{E}_{s+1} V'_s - \pi (a_{s+1}, c) + K (1 - \beta) \right] \quad \text{if } c \leq c^U_s$$

$$= \lambda_{s+1} \left[ \frac{\lambda_{s+1} \pi (a_{s+1}, c) + K (1 - \beta)}{1 - \beta \lambda_{s+1} (1 + \beta \lambda_{s+1})} \right]$$

We can then show by contradiction that $c^U_s < c^D_s$. Suppose instead that $c^U_s \geq c^D_s$ so the marginal deterministic firm has non-positive option value of waiting at $s$ under uncertainty, i.e. $V_s (c^D) \leq 0$. By definition $\pi (a_s, c^D_s) = K (1 - \beta)$ and so $V_s (c^D) = \max \left\{ 0, \beta \mathbb{E}_s V'_s (c^D_s) \right\}$. Moreover, $\pi (a_{s+1}, c^D_s) < K (1 - \beta)$ when $\tau_s < \tau_{s+1}$, which implies that $\mathbb{E}_s V'_s > 0$ and therefore $V_s (c^D) > 0$. This contradiction implies that $c^U_s < c^D_s$.

The marginal firm at $s$ under uncertainty has $V_s (c^U_s) = \max \left\{ 0, \beta \mathbb{E}_s V'_s (c^U_s) \right\} - \pi (a_s, c^U_s) + K (1 - \beta)$ and we can solve for $c^U_s$ by equating the second term in curly brackets to zero and simplifying to obtain

$$\pi (a_s, c^U_s) + \beta \lambda_{s+1} \frac{\pi (a_{s+1}, c^U_s)}{1 - \beta \lambda_{s+1}} = K (1 - \beta) \left( 1 + \frac{\beta \lambda_{s+1}}{1 - \beta \lambda_{s+1}} \right)$$

Starting at $s = 1$, replacing $\pi$ with its value in (1) and simplifying we obtain the cutoff expression (7) in the text

$$a_1 (c^U_1)^{1 - \sigma} + \frac{\beta \lambda_{12}}{1 - \beta \lambda_{22}} a_2 (c^U_1)^{1 - \sigma} = K (1 - \beta) \left( 1 + \frac{\beta \lambda_{12}}{1 - \beta \lambda_{22}} \right)$$

$$a_1 (c^U_1)^{1 - \sigma} \left( 1 + u(\gamma) \frac{a_2}{a_1} \right) = K (1 - \beta) (1 + u(\gamma))$$

$$c^U_1 = \left( \frac{a_1}{K (1 - \beta)} \right)^{\frac{1}{1 - \sigma}} \times \left( \frac{1 + u(\gamma)}{1 + u(\gamma)} \right)^{\frac{1}{1 - \sigma}} = c^D_1 	imes U_1 (\omega, \gamma)$$

The last line uses the expressions given in the main text: $\omega \equiv a_2/a_1 = (\tau_2/\tau_1)^{-\sigma}$, $u(\gamma) \equiv \frac{\beta \gamma \lambda_2}{1 - \beta \lambda_{22}}$, $\gamma \equiv 1 - \lambda_{11}$, $\gamma \lambda_2 = \lambda_{12}$.

Proof of Prop 1(b): $c^U_1 < c^D_1$ and $dc^U_1/d\gamma > dU_1/d\gamma < 0$ all $\gamma$ iff tariff increases are possible

Since $c^U_1/c^D_1 = U_1$ we must show $U_1 < 1$ iff tariff increases are possible. From the definition in (8) we obtain $U_1 < 1$ iff $u(\gamma) \omega < u(\gamma)$, which is true iff $T_2 > T_1$ and $\lambda_{12} > 0$ so that $u(\gamma) > 0$.

Since $c^U_1 = c^U_1 U_1 (\omega, \gamma)$ (part a) we can use (8) to obtain

$$\gamma \frac{d \ln c^U_1}{d \gamma} = \gamma \frac{d \ln U_1 (\omega, \gamma)}{d \gamma} = \frac{1}{\sigma - 1} \frac{u - 1}{1 + u - uw} < 0$$

(36)
where the inequality holds only if $\omega < 1 \iff \tau_2 > \tau_1$ and $u(\gamma) > 0$.

Proof of Prop 1(c): $c_1^U < c_1^D = c_0^D$

If $\tau_2 > \tau_1$ and $u > 0$ then $c_1^U < c_1^D$, (part b). Since we assume that $\lambda_{00} = 1$ we have $c_0^D = c_0^U$. If $\tau_0 \leq \tau_1$ then $c_1^D \leq c_0^D$ (from (3)) and therefore $c_1^U < c_1^D \leq c_0^D = c_0^U$. If tariff increases are not possible ($\tau_2 = \tau_1$ and/or $u = 0$) then $c_1^U = c_1^D = c_0^D$ if $\tau_0 < \tau_1$. The first equality is shown in part (b) and the inequality is from (3).

QED

A.2 Proof of Proposition 2

Definition D1: Properties of conditional distribution of $a_s, \hat{\Lambda}(a_s, \gamma)$

1) $a_s \equiv (\tau_m \sigma)^{-\sigma} ((\sigma - 1) P_s / d)^{1-\sigma} E$ so $\hat{\Lambda}(a_s, \gamma)$ reflects the exogenous distribution $\Lambda(\tau_m, \gamma)$ and the endogenous adjustment of $P_s$.

2) $\Lambda(\tau_m, \gamma)$ is the time invariant conditional distribution of policy and we assume that $\Lambda(\tau_{m+1}, \gamma)$ FOSD $\Lambda(\tau_m, \gamma)$ for $m = 0, 1$ and any $\tau_2 > \tau_1 \geq \tau_0$.

3) The number of states in this case is $s = 0, 1, 2T$ where $T$ is the number of periods since the policy increased to $\tau_2$ and is finite.

4) In section 2.3 we show $a_0 > a_1 > a_{2T}$ and $a_{2T} > a_{2T-1}$, which is also the ordering under uncertainty around the deterministic setting.

Definition D2: Positive state transition probabilities in $\hat{\Lambda}(a_s, \gamma)$

1) $\lambda_{ss'} > 0$ for all $s = 0, 1$ and $s' = 0, 1, 2T$ for $T = 0$, (2) $\lambda_{ss'} = 0$ for all $s \rightarrow s'$ where $s = 0, 1$ and $s' = 2T$ for $T > 0$, (3) $1 - \lambda_{ss} > 0$ for all $s = 0, 1, 2T$, (4) $\lambda_{22} > 0$ for all $s = 2T \rightarrow s' = 2T + 1$, (5) 0 otherwise.

Lemma 2: If $\Lambda(\tau_{m+1}, \gamma)$ FOSD $\Lambda(\tau_m, \gamma)$ for $m = 0, 1$ and any $\tau_2 > \tau_1 \geq \tau_0$ then $\hat{\Lambda}(a_s, \gamma)$ inherits the stochastic dominance properties of $\Lambda(\tau_m, \gamma)$ that are sufficient for a single entry threshold:

1) $\hat{\Lambda}(a_0, \gamma)$ FOSD $\hat{\Lambda}(a_s, \gamma)$ for all $s \neq 0$

2) $\hat{\Lambda}(a_1, \gamma)$ FOSD $\hat{\Lambda}(a_{2T}, \gamma)$ for all $T$

3) $\hat{\Lambda}(a_{2T}, \gamma)$ FOSD $\hat{\Lambda}(a_{2T-1}, \gamma)$ for all $T > 0$

Proof of Lemma 2: By D1.1 $a$ is inversely related to $\tau$ and by D1.4 $a_0 > a_1 > a_{2T}$. By D1.2 $\Lambda(\tau_{m+1}, \gamma)$ FOSD $\Lambda(\tau_m, \gamma)$ for $m = 0, 1$ and any $\tau_2 > \tau_1 \geq \tau_0$ and this is equivalent to $\hat{\Lambda}(a_s, \gamma)$ FOSD $\hat{\Lambda}(a_{s'}, \gamma)$ for $s = 0, 1$ and $s' = 1, 2T$. The positive transition probabilities of $\hat{\Lambda}(a_s, \gamma)$ in D2 are the same as $\Lambda(\tau_m, \gamma)$ This implies $\hat{\Lambda}(a_s, \gamma)$ FOSD $\hat{\Lambda}(a_{s'}, \gamma)$ for $s = 0, 1$ and $s' = 1, 2T$. This proves parts (1) and (2). Part (3) follows because the only difference between $\hat{\Lambda}(a_{2T-1}, \gamma)$ and $\hat{\Lambda}(a_{2T}, \gamma)$ is that in the latter all probability weight from $a_{2T-1}$ is shifted to a higher value, $a_{2T}$ (D1.4).

QED

Lemma 3: (Entry threshold, GE): For any given policy regime $\Lambda(\tau_m, \gamma)$ that exhibits uncertainty persistence, and each firm $c$, there is a unique threshold $a^U_c(\gamma, c)$ per state above which a firm enters into exporting.

Proof of Lemma 3: We rewrite (2) recursively and substitute in (4) to obtain (62), as we did in Lemma 1:

$$V_s = \max \{0, \beta E_s V_{s'} - \pi(a_s, c) + K (1 - \beta)\}$$

where the option value of waiting is $V_s = \Pi(a_s, c, \gamma) - \Pi_c(a_s, c, \gamma) + K$ and $E_s V_{s'} \equiv E_s [\Pi(a'_s, c, \gamma) - \Pi_c(a'_s, c, \gamma) + K].$

Now we focus on cutoffs as functions of $a$ instead of $\tau$. By Lemma 2 and D2 above we have properties analogous to Lemma 1:

1) The ex-ante option value of waiting requires the expectation over $a$, so $E_s V_{s'} = \int V_{s'} d\Lambda(a_s, \gamma)$.

2) Because $-\pi(a_s, c)$ is decreasing in $a$ it is more attractive to wait at lower $a$ since the second element of (62) and therefore $V_s$ would be higher, all else equal.

3) Since $\hat{\Lambda}(a_s, \gamma)$ exhibits uncertainty persistence (from Lemma 2), $\int V_s d\Lambda(a_s, \gamma)$ is decreasing in the current $a_s$ if $V_s$ is decreasing in $a$.

Given (3) if we start with a decreasing $V_s$ then the fixed point to this iteration is also decreasing in $a$. By properties (2) and (3), $\beta E_s V_{s'} - \pi(a_s, c)$ is decreasing in $a$, so there is some $a^U_c(\gamma, c)$ above which the firm
value is higher if exporting and below which the opposite is true. \[ \text{QED} \]

**Proof Prop. 2(a):** unique entry cutoff \( c^U_1 = c^D_1 U_1 (\tilde{\omega}_V, \gamma) P_1 / P^D_1 \)

Lemma 3 shows that for each firm \( v \) there is a single cutoff \( c^U_v (\gamma, c_v) \) above which they enter. All firms face the same \( a \) and \( \gamma \) and have different \( c \) so for a given \( a \) and \( \gamma \) there is a single cutoff \( c^U_1 \) and only firms with costs below it will enter. Using the definitions of \( c^D_1 \), \( a_1 \), and \( U_1 \) we must show

\[
c^U_1 = \left[ \frac{a_1}{(1 - \beta) K} \right] \pi^+ \left[ \frac{1 + u (\gamma) \omega g}{1 + u (\gamma)} \right] \pi^+ \tag{37}
\]

where \( g \) is defined in (14). The proof is identical to part (a) of proposition 1 except now (34) reflects the transition dynamics in \( P \) after the tariff increases so we have

\[
E_{s=2T} V'_{s} = \lambda_{22} [\beta E_{s=2T+1} V'_{s} - \pi (a_{s=2T+1}, c) + K (1 - \beta)] \quad \text{if } c \leq c^U_1 \tag{38}
\]

We can solve this forward to obtain \( E_{s=1} V'_{s} = -\lambda_{22} \sum_{t=0}^{\infty} (\beta \lambda_{22})^t \pi (a_{s=2t+1}, c) + \frac{\lambda_{22}}{1 - \beta \lambda_{22}} K (1 - \beta) \). Replacing this in (33) and simplifying we obtain

\[
E_{s=1} V'_{s} = \frac{\lambda_{12}}{1 - \beta \lambda_{22}} \left[ - (1 - \beta \lambda_{22}) \sum_{t=0}^{\infty} (\beta \lambda_{22})^t \pi (a_{s=2t+1}, c) + K (1 - \beta) \right] \tag{39}
\]

Finally, we find the cutoff expression for the marginal firm in \( s = 1 \), the one with \( V_{1} (c^U_1) = 0 \). So we must solve for \( c^U_1 \) in the expression below, as in proposition 1, but now using the value of \( E_{s=1} V'_s \) derived in (39).

\[
\beta E_{s=1} V'_s (c^U_1) - \pi (a_1, c^U_1) + K (1 - \beta) = 0
\]

After simplification we obtain (37).

**Proof Prop. 2(b):** when tariff increases are possible then \( c^U_1 < c^D_1 \) and \( P_1 > P^D_1 \)

From part (a) \( c^U_1 = c^D_1 U_1 (\tilde{\omega}_V, \gamma) P_1 / P^D_1 \) so \( c^U_1 < c^D_1 \) iff \( U_1 (\tilde{\omega}_V, \gamma) P_1 / P^D_1 < 1 \). From its definition we see that \( U_1 (\tilde{\omega}_V, \gamma) < 1 \) if \( u > 0 \) and \( \tilde{\omega} = \left( \frac{\omega}{\tau_1} \right) > 1 \). Note that \( \tilde{\omega} \) measures the ratio of the average operating profits under the worst case scenario relative to state 1, so \( \tilde{\omega} < 1 \) if \( \tau_2 > \tau_1 \). We can now use \( U_1 < 1 \) to prove the main result by contradiction. Assume first that \( P_1 \leq P^D_1 \). In this case \( c^U_1 < c^D_1 \) because \( U_1 < 1 \). But if \( c^U_1 < c^D_1 \) then \( P_1 > P^D_1 \) because under uncertainty there are fewer varieties exported but all else is unchanged. Next, assume \( c^U_1 / c^D_1 \geq 1 \). In this case \( P_1 / P^D_1 \leq 1 \) because under uncertainty there would be more varieties exported but all else is unchanged. But this violates the equilibrium condition derived in (a) since in order for \( c^U_1 / c^D_1 \geq 1 \) we require \( U_1 (\tilde{\omega}_V, \gamma) P_1 / P^D_1 \geq 1 \), which is not possible if \( P_1 / P^D_1 \leq 1 \) and \( U_1 < 1 \).

**Proof Prop. 2(c):** when tariff increases are possible then \( P_1 > P_0 = P^D_0 \) and \( c^U_1 < c^D_1 = c_0^D \)

In part (b) we show \( P_1 > P^D_1 \) (and \( c^U_1 < c^D_1 \)) so it is sufficient to show \( P^D_1 \geq P^D_0 = P_0^D \) (and \( c^U_1 \leq c^D_1 = c_0^D \)) which we did in section 2.6.3 in (10). \[ \text{QED} \]

### A.3 Proof of Proposition 3

**Proof Prop. 3(a)** Workers per period indirect utility in state \( s \) is \( \tilde{\mu} P_{s}^{-\mu} \) where \( \tilde{\mu} = \omega \mu \mu^2 (1 - \mu)^{(1 - \mu)} \) is constant. Eliminating MFN policy uncertainty lowers the price index (proposition 2) and thus increases consumer welfare in the MFN state.

**Proof Prop. 3(b)** Let \( W_s \) denote the expected welfare of a consumer starting in \( s \). We first solve for the equilibrium value of \( W_1 \) and then show it is lower than under an agreement without policy uncertainty, \( W_1 < W_0 = \tilde{\mu} (P^D_0)^{-\mu} / (1 - \beta) \). In general recursive form we can write \( W_s = \tilde{\mu} P_{s}^{-\mu} + \beta E_s W'_s \) for \( s = 0, 1, 2T \).
Using the policy matrix we can write

\[ W_1 = \mu P_1^{-\gamma} + \beta [\lambda_{11} W_1 + \lambda_{10} W_0 + \lambda_{12} W_2] \]
\[ W_{2t} = \mu P_{2t}^{-\gamma} + \beta \lambda_{22} W_{2t+1} + \beta (1 - \lambda_{22}) W_1 \quad \text{all } t \geq 0 \]

Solving \( W_{2t} \) forward we obtain

\[ W_{20} (1 - \beta \lambda_{22}) = (1 - \beta \lambda_{22}) \sum_{t=0}^{\infty} \left( \beta \lambda_{22} \right)^t \mu P_{2t}^{-\gamma} + \beta (1 - \lambda_{22}) W_1 \]

This allows us to obtain \( W_1 \) by solving the following 3-equation linear system

\[ W_m (1 - \beta \lambda_{mm}) - \beta \left( \Sigma_{m\neq m'} \lambda_{mm'} W_{m'} \right) = w_m \quad \text{for } m = 0, 1, 2 \tag{40} \]

where \( w_m = \mathbb{E}_{s \in m} \tilde{\mu} P_s^{-\gamma} \) is the average period welfare while under policy \( m \). It is clear that \( W_1 \) is weighted average of \( w_m \), as given in (19) and by solving the system we obtain the value of those weights listed below. These weights represent the expected number of periods that a consumer in \( m = 1 \) expects to spend in each policy state. This intuitive result is a property of the type of Markov chain in this setting, as we show in the Online Appendix C.5.

\[ n_1 = \left(1 - \beta \left(1 - \gamma + \lambda_{21} \frac{\gamma \lambda_2 \beta}{1 - \beta \lambda_{22}}\right)\right)^{-1} ; \quad n_0 = \gamma (1 - \lambda_2) \frac{\beta}{1 - \beta} n_1 ; \quad n_2 = \gamma \lambda_2 \frac{\beta}{1 - \beta} n_1 \tag{41} \]

It is then straightforward to show \( W_1 (\gamma) = \sum_{m=0,1,2} n_m (\gamma) w_m (\gamma) < w_0 / (1 - \beta) = W_0 \) since \( P_0 < P_1 < P_2 \) (Proposition 2) implies \( w_0 > w_1 > w_2 \) and the agreement state has the highest weight on \( w_0 \).

**Proof Prop. 3(c)** A first order log approximation of (19) around \( \gamma = 0 \) yields

\[
\ln \frac{W_1 (\gamma > 0)}{W_1 (\gamma = 0)} \approx \gamma \frac{d \ln w_1}{d \gamma} \bigg|_{\gamma=0} + \gamma (1 - \beta) \sum_{m=0,2} \frac{dn_m}{d \gamma} \bigg|_{\gamma=0} \times \ln \frac{w_m}{w_1} \bigg|_{\gamma=0} = \gamma \frac{d \ln w_1}{d \gamma} \bigg|_{\gamma=0} + \sum_{m=0,2} \frac{n_m}{n_1} \times \ln \frac{w_m}{w_1} \bigg|_{\gamma=0} = \frac{d \ln w_1}{d \gamma} \bigg|_{\gamma=0} + \sum_{m=0,2} \frac{\beta \lambda_{1m}}{1 - \beta \lambda_{mm}} \times \ln \frac{w_m}{w_1} \bigg|_{\gamma=0} = -\mu \gamma \frac{d \ln P_1}{d \gamma} \bigg|_{\gamma=0} + \sum_{m=0,2} \frac{\beta \lambda_{1m}}{1 - \beta \lambda_{mm}} \times \ln \mathbb{E}_{s \in m} \left( \frac{P_{sD}}{P_s} \right)^{1-\mu} \]

The first line uses \( n_1|_{\gamma=0} = (1 - \beta)^{-1} \) and \( n_m|_{\gamma=0} = 0 \) otherwise and \( \frac{w_m}{w_1}|_{\gamma=0} = 1 \approx \ln \frac{w_m}{w_1} |_{\gamma=0} \). The second uses \( \gamma (1 - \beta) \frac{dn_m}{d \gamma} \bigg|_{\gamma=0} = \frac{n_m}{n_1} \) for \( m = 0, 2 \). The third line uses \( n_m \) from (41), and the last one uses \( w_m (P) = \mathbb{E}_{s \in m} \tilde{\mu} P_s^{-\gamma} \).

**QED**

### A.4 Multi-industry price index aggregation

Whenever economic conditions today are at least as good as the past, \( c^V_t \geq \max \{c^V_{t-T} \}_{T=0} \) for each \( V \), we can write the price index as a function of the vector of current tariffs and cutoffs \( (\tau_1, c_1) \).

\[ P_t (\tau_1, c_1) = \left[ \sum_{V \in \mathcal{V}} NV \int_0^{c^V_t} (w \tau_1 d V c / \rho)^{1-\sigma} d G_V (c) + \sum_{V \neq C \in \mathcal{V}} \int_{v \in \mathcal{V}_{C \neq v}} (c_v / \rho)^{1-\sigma} dv \right]^{-1} \tag{42} \]
The deterministic policy entry cutoff is still defined by the expression in (3). As we show in the working paper, in a multi-industry model there is a cutoff for each industry \( V \) that depends on the aggregate price index, \( c_V (P^D, \tau_V) \). The elasticity of entry with respect to tariffs now requires comparative statics on a system of equations that determine the foreign exporter entry cutoffs in each of the \( V \) industries and one equation for the domestic price index. To verify that this system has a unique equilibrium we first make use of the fact that the cutoffs are linear in \( P \) and in constant parameters. So any industry cutoff can be written as a linear function of some base industry cutoff, \( c_{sb}^D \), and relative parameters, i.e. \( c_{sb}^D = c_{sb}^D \kappa_{Vb} \)

where \( \kappa_{Vb} \equiv \left[ \frac{\left( \tau_{sb} / \tau_{sb} \right) - \sigma}{K_V / K_b} \right] \frac{1}{\kappa_V} \frac{dV}{sb} \). Using this we write the reduced form index as

\[
P \left( \tau_s, c_{sb}, c_{sb}^D \right) (c_{sb}^D \kappa_{Vb}) \)

which is a positive function that is continuous and non-increasing in \( c_{sb}^D \). The entry schedule for the base industry has positive slope, since \( \partial c_{sb}^D / \partial P > 0 \) and \( c_{sb}^D |_{P \to 0} = 0 \). Therefore these two schedules intersect exactly once at an equilibrium.

We can define an implicit solution to the system of \( V + 1 \) equations \( P (c^D, \tau) \) and \( c_V (P^D, \tau_V) \) for each \( V \) so the total change due to the tariff can be found by totally differentiating (43)

\[
dln P (c^D, \tau) = \sum_V \frac{\partial \ln P (c^D, \tau)}{\partial \ln c^D} dln c^D + \sum_V \frac{\partial \ln P (c^D, \tau)}{\partial \ln \tau_V} dln \tau_V
\]

\[
= \frac{k - \sigma + 1}{(1 - \sigma)} I \sum_V \frac{\tau_V R_V}{\sum \tau_V R_V} dln c^D + I \sum_V \frac{\tau_V R_V}{\sum \tau_V R_V} dln \tau_V
\]

where \( I \equiv \sum_V \tau_V R_V / E \) is the share of industry \( V \) expenditure. We derive the two partial elasticities used in the second line in the Online Appendix C.1.1

\[B \quad \textbf{Data and Estimation Appendix}\]

\[B.1 \quad \textbf{Data sources and definitions}\]

- **Change in Advalorem Tariffs \( \Delta \tau_V \):** Log change in 1 plus the statutory advalorem MFN tariff rate aggregated to the HS6 level between 2005 and 2000. Source: TRAINS via WITS.
- **Change in AVE Tariffs \( \Delta \tau_V \):** Log change in 1 plus the advalorem equivalent (AVE) of the MFN tariff rate at the HS6 level between 2005 and 2000. For specific tariffs, the AVE is given by the ratio of unit duty to the average 1996 import unit value. Source: TRAINS for tariff rates and COMTRADE for unit values via WITS.
- **Column 2 Tariff \( \tau_2V \):** Log of 1 plus the column 2 (Smoot-Hawley) tariff rate at the HS6 level. For specific tariffs at the HS8, base year unit values from 1996 used for all years to compute the AVE tariff and then average at the HS6 level. Source: TRAINS for tariff rates and COMTRADE for unit values via WITS.
- **Pre-WTO Uncertainty:** Measure of uncertainty from the model 1 – \( \left( \frac{\tau_V}{\tau_1V} \right) ^{-\sigma} \) computed using year 2000 column 2 and MFN tariff rates.
- **Change in Transport Costs \( \Delta D_V \):** Log change in the ratio of trade values inclusive of costs, insurance and freight (CIF) to free on board value (FOB). Source: CIF/FOB ratios constructed at HS6 level using disaggregated data from NBER.
- **Change in TTBs:** Indicators for temporary trade barriers in-force including anti-dumping duties, countervailing duties, special safeguards, and China-specific special safeguards. Data are aggregated up to HS6 level. Source: World Bank Temporary Trade Barriers Database (Bown, 2012).
- **Change in MFA:** Indicators for in-force Multi-Fiber Agreement on Textiles and Clothing (MFA/ATC) quotas aggregated to the HS6 level and concorded through time. Source: Brambilla et al. (2010)

---

\[69\text{It can be shown that } \partial P / \partial \kappa_{Vb}^D \leq 0 \text{ for all } V, \text{ strictly so for small enough } c, \text{ and } \partial c_{sb}^D / \partial \kappa_{Vb} = \kappa_{Vb} \text{ for all } V \neq b. \text{ Continuity holds provided that the distribution of firms in each industry is not bounded above so there is always at least one active exporter.}\]
• Change in No. of HS-10 Traded Products: Change in log count of traded HS10 products within each HS6 industry from 2000 to 2005. Source: disaggregated data from NBER

B.2 Expenditure share, import penetration and risk counterfactuals

Import penetration in manufacturing is Chinese imports over U.S. expenditure on manufacturing, $M_{Ch,t}/E_t$. We define total manufacturing expenditure, given by $E = \mu L$ in the model, as total manufacturing shipments less net manufacturing exports, $E_t = \text{Manuf. Shipments}_t - \text{Exports}_t + \text{Imports}_t$. We compute $\mu = 0.86$ as the share of manufacturing in total expenditure ($=\text{Gross Output} - \text{Total Net Exports}$) in 2005.

For each year from 1990 to 2010 we obtain manufacturing shipments from the U.S. Census Bureau and manufacturing exports and imports from the USITC. We include tariffs and transport costs in total imports, as our model requires. To compute the counterfactual imports if uncertainty were reintroduced, we use the Chinese import penetration in the data from each year for 2002-2010 to recompute the upper bound on the price effect $\bar{g}$, ranging from 1.05 in 2002 up to 1.11 in 2010 and the change in price index from uncertainty $\gamma_d \ln P_1(\bar{\tau}, \bar{D})$ from equation (27) for each year. We use this to compute the counterfactual imports from China normalized by expenditure, $M^{CF}_{Ch,t}/E_t$, which we plot in Figure 1.

To find the share of average import growth from a pure risk reduction, we compute import growth from reducing uncertainty as if the tariffs were at the long run mean. We adjust imports to levels implied by the model at the mean tariff. The procedure uses the 2005 import penetration to compute the price elasticity to a tariff change. We then compute the change in imports implied by a partial equilibrium tariff change to the mean for each industry adjusted by the new, higher price index. We compute import penetration and the weight of each industry in the price index at the mean tariff using these counterfactual import levels. We obtain a new upper bound on the price index and the partial effect through $\bar{g}$. With the model quantities all adjusted to their levels at the mean of the tariff distribution, we can then compute $\frac{d\ln P_1}{d\gamma}$ and the full GE effect on exports around the mean.
Figures and Tables

Figure 1: Chinese Import Penetration—Actual vs. Counterfactual under Policy Uncertainty.

![Graph showing Chinese Import Penetration Actual vs. Counterfactual under Policy Uncertainty]

Notes: Actual import penetration ratio defined as manufacturing imports from China as share of total expenditure on manufacturing (total shipments - net exports). Counterfactual line adjusts Chinese imports downward as if uncertainty reintroduced in any year after 2001. See data appendix for further details.

Figure 2: Partial vs. general equilibrium effects of an unanticipated tariff decrease.

![Diagram showing partial vs. general equilibrium effects]

Notes: I: initial equilibrium. PE: equilibrium after tariff reduction in partial equilibrium. GE: general equilibrium.

Figure 3: Transition dynamics after an unanticipated tariff increase.

![Diagram showing transition dynamics]

Notes: 1\(^D\): initial MFN deterministic equilibrium, 2\(^{TR}\): lower bound cutoff and price after tariff increase, 2\(^{TR} \rightarrow 2^{D}\): transition path after switch to column 2 until steady state, 1\(^U\): equilibrium under uncertainty, state 1 (MFN).
Figure 4: China’s Export Growth 2000-2005 vs. US pre-WTO tariff threat — non-parametric and linear fit.

Notes: Linear fit from OLS regression: export growth=1.05+0.92×ln(τ_2/τ_{MFN}) where τ_2 and τ_{MFN} are the column 2 and MFN tariff factors in 2000; both coefficients are significant at the 1% level. The non-parametric fit uses a running-line least-squares smoothing (lowess).

Figure 5: China’s Export Growth 2000-2005 vs. US Pre-WTO Policy Uncertainty Measure — semi-parametric and linear fit

Notes: Both fits regress export growth on changes in transport costs, tariffs and section dummies. The linear fit uses OLS and also includes −(τ_2/τ_{MFN})^{-3}, which the semi-parametric uses as an argument of the local polynomial estimated using the Robinson(1988) semi-parametric estimator. We plot the fit against 1 − (τ_2/τ_{MFN})^{-3} for ease of comparison with the uncertainty variable used in the baseline. The plot without sector dummies is similar.
### Table 1: Summary statistics by pre-WTO policy uncertainty

<table>
<thead>
<tr>
<th>Uncertainty</th>
<th>Low</th>
<th>High</th>
<th>Total</th>
</tr>
</thead>
<tbody>
<tr>
<td>Chinese export growth to US (Δln, 2005-2000)</td>
<td>1.18</td>
<td>1.36***</td>
<td>1.29</td>
</tr>
<tr>
<td></td>
<td>[1.788]</td>
<td>[1.603]</td>
<td>[1.672]</td>
</tr>
<tr>
<td>MFN tariff (ln), 2000</td>
<td>0.028</td>
<td>0.044</td>
<td>0.039</td>
</tr>
<tr>
<td></td>
<td>[0.036]</td>
<td>[0.048]</td>
<td>[0.045]</td>
</tr>
<tr>
<td>Column 2 tariff (ln), 2000</td>
<td>0.158</td>
<td>0.393</td>
<td>0.311</td>
</tr>
<tr>
<td></td>
<td>[0.096]</td>
<td>[0.116]</td>
<td>[0.156]</td>
</tr>
<tr>
<td>Ratio of Col 2 to MFN tariff</td>
<td>1.146</td>
<td>1.428</td>
<td>1.33</td>
</tr>
<tr>
<td></td>
<td>[0.090]</td>
<td>[0.140]</td>
<td>[0.183]</td>
</tr>
<tr>
<td>Potential profit loss if MFN revoked (pre WTO)</td>
<td>0.303</td>
<td>0.636</td>
<td>0.52</td>
</tr>
<tr>
<td></td>
<td>[0.175]</td>
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<tr>
<td>Observations</td>
<td>1132</td>
<td>2110</td>
<td>3242</td>
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</table>

Notes:
- Means with standard deviations in brackets. Low: subsample of industries in the bottom tercile of pre-WTO uncertainty (ranked by $\tau_2$); High refers to the rest of the sample. Total includes the full sample used in baseline Table 2. *** 1% significance level for difference of mean export growth between low and high subsamples.

### Table 2: Export Growth from China (2000-2005)

<table>
<thead>
<tr>
<th>Uncertainty Pre-WTO</th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
</tr>
</thead>
<tbody>
<tr>
<td>[*]</td>
<td>0.682***</td>
<td>0.731***</td>
<td>0.687***</td>
<td>0.703***</td>
</tr>
<tr>
<td>[-]</td>
<td>[0.158]</td>
<td>[0.154]</td>
<td>[0.186]</td>
<td>[0.185]</td>
</tr>
<tr>
<td>Change in Tariff (Δln)</td>
<td>-9.702***</td>
<td>-3.969***</td>
<td>-6.608</td>
<td>-3.894***</td>
</tr>
<tr>
<td>[-]</td>
<td>[4.473]</td>
<td>[0.702]</td>
<td>[5.057]</td>
<td>[0.704]</td>
</tr>
<tr>
<td>Change in Transport Costs (Δln)</td>
<td>-2.556***</td>
<td>-2.646***</td>
<td>-2.562***</td>
<td>-2.596***</td>
</tr>
<tr>
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<td>[0.474]</td>
<td>[0.468]</td>
<td>[0.474]</td>
<td>[0.469]</td>
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<tr>
<td>Constant</td>
<td>0.895***</td>
<td>0.887***</td>
<td>[0.0881]</td>
<td>[0.0877]</td>
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</table>

Notes:
- Robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.10. Predicted sign of coefficient in brackets under variable. Uncertainty measure uses U.S. MFN and Column 2 Tariffs to construct profit loss measure at $\sigma=3$. All specifications employ OLS and 2 and 4 impose theoretical constraint on tariffs and transport cost coefficients: $b_\tau=b_\sigma(\sigma/(\sigma-1))$. 

- R-squared 0.03 n/a 0.05 n/a
- Sector fixed effects no no yes yes
- Restriction p-value (F-test) 0.195 1 0.588 1
Table 3: Export Growth from China (2000-2005): Robustness to NTBs

<table>
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<tr>
<th>Specification:</th>
<th>Baseline</th>
<th>+MFA/TTB</th>
<th>+Section FE</th>
<th>IV (NTB)</th>
<th>Constrained</th>
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<tr>
<td>Uncertainty Pre-WTO</td>
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<td>0.688***</td>
<td>0.694***</td>
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<td>[0.186]</td>
<td>[0.185]</td>
<td>[0.184]</td>
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<tr>
<td>Change in Tariff (Δln)</td>
<td>-9.702**</td>
<td>-8.791*</td>
<td>-7.47</td>
<td>-7.61</td>
<td>-3.948***</td>
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<td>[4.546]</td>
<td>[5.057]</td>
<td>[5.046]</td>
<td>[0.700]</td>
</tr>
<tr>
<td>Change in Transport cost (Δln)</td>
<td>-2.556***</td>
<td>-2.548***</td>
<td>-2.588***</td>
<td>-2.596***</td>
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<td>Change in MFA quota status</td>
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<td>-0.311**</td>
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<tr>
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<tr>
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<td>Constant</td>
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<td>0.912***</td>
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<td>[0.0881]</td>
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<tr>
<td>R-squared</td>
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<td>Sector fixed effects</td>
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<td>0.482</td>
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</table>

Notes: Robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.10. Predicted sign of coefficient in brackets under variable. Specifications 1-3 employ OLS and 5 impose theoretical constraint on tariffs and transport cost coefficients: $b^\tau=bd^\tau/\sigma/(\sigma-1)$. Specification 4 employs IV. Excluded instruments for Change in NTB are NTB indicators for 1998 and 1997. Uncertainty measure uses U.S. MFN and Column 2 Tariffs to construct profit loss measure at $\sigma=3$.

Table 4: Export Growth from China: Non-linear and linear estimates

<table>
<thead>
<tr>
<th></th>
<th>1</th>
<th>2</th>
<th>3</th>
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</thead>
<tbody>
<tr>
<td>estimation method</td>
<td>NLLS</td>
<td>OLS</td>
<td>NLLS</td>
<td>OLS</td>
</tr>
<tr>
<td>Uncertainty (pre-WTO)</td>
<td>0.823***</td>
<td>0.646***</td>
<td>0.668**</td>
<td>0.542***</td>
</tr>
<tr>
<td>[-]</td>
<td>[0.305]</td>
<td>[0.149]</td>
<td>[0.338]</td>
<td>[0.184]</td>
</tr>
<tr>
<td>Change in Tariff (ln)</td>
<td>-6.594***</td>
<td>-6.376***</td>
<td>-6.302***</td>
<td>-6.186***</td>
</tr>
<tr>
<td>[-]</td>
<td>[1.242]</td>
<td>[1.246]</td>
<td>[1.248]</td>
<td>[1.249]</td>
</tr>
<tr>
<td>Change in Transport Costs (ln)</td>
<td>-4.396***</td>
<td>-4.251***</td>
<td>-4.202***</td>
<td>-4.124***</td>
</tr>
<tr>
<td>[-]</td>
<td>[0.828]</td>
<td>[0.831]</td>
<td>[0.832]</td>
<td>[0.833]</td>
</tr>
<tr>
<td>Constant</td>
<td>1.579***</td>
<td>0.908***</td>
<td>1.016</td>
<td>0.0844</td>
</tr>
</tbody>
</table>

Upper bound on average export growth from removing uncertainty

<table>
<thead>
<tr>
<th></th>
<th>0.29</th>
<th>0.34</th>
<th>0.25</th>
<th>0.28</th>
</tr>
</thead>
<tbody>
<tr>
<td>from removing uncertainty</td>
<td>[0.0705]</td>
<td>[0.0782]</td>
<td>[0.0867]</td>
<td>[0.0964]</td>
</tr>
</tbody>
</table>
Table 5: Contribution of Policy Uncertainty to Export and Variety Growth (2000-2005): NLLS estimates

<table>
<thead>
<tr>
<th></th>
<th>No GE</th>
<th>Partial GE</th>
<th>Full GE</th>
</tr>
</thead>
<tbody>
<tr>
<td>Average export growth from lower uncertainty (Δln)</td>
<td>0.29</td>
<td>0.26</td>
<td>0.22</td>
</tr>
<tr>
<td>Average variety growth from lower uncertainty (Δln)</td>
<td>0.53</td>
<td>0.47</td>
<td>0.43</td>
</tr>
</tbody>
</table>

Notes:
Employs NLLS estimates (column 1 Table 4). Each column represents the difference between average predicted change with and without uncertainty effect. The No GE column assumes a small exporter so the price index is unchanged today and under higher tariffs. The partial GE assumes no uncertainty effect on contemporaneous price index but allows for an effect if tariffs increase substantially. Full GE allows for contemporaneous and future price effects. See text for formulas.

Table 6: Trade Cost Equivalents of Trade Policy Uncertainty Change

<table>
<thead>
<tr>
<th></th>
<th>Chinese Export Value [-]</th>
<th>Chinese Export Entry [-]</th>
<th>US Price index and firms' domestic sales [+ ]</th>
<th>US Consumer Welfare [-]</th>
</tr>
</thead>
<tbody>
<tr>
<td>Ad Valorem tariff equivalent of TPU (Δln)</td>
<td>0.036</td>
<td>0.071</td>
<td>0.076</td>
<td>0.076</td>
</tr>
<tr>
<td>Ad Valorem transport cost equivalent of TPU (Δln)</td>
<td>0.054</td>
<td>0.11</td>
<td>0.096</td>
<td>0.096</td>
</tr>
</tbody>
</table>

Notes:
Quantification uses NLLS estimates (column 1, Table 4). See text for details of calculation. Equivalents for outcome x in any period after uncertainty changes but applied policies remain unchanged after any transition dynamics in the case of increased uncertainty. [-/+] denotes sign of derivative of column variables wrt trade cost.

Table 7: Risk Decomposition of Counterfactual Uncertainty Effects

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Fraction of uncertainty effect due to:</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Risk reduction (at LR expected tariffs)</td>
<td>0.82</td>
<td>0.79</td>
<td>0.43</td>
<td>0.43</td>
</tr>
<tr>
<td>Expected tariff reduction</td>
<td>0.18</td>
<td>0.21</td>
<td>0.57</td>
<td>0.57</td>
</tr>
<tr>
<td>Full GE Effect (at MFN tariffs, Δln)</td>
<td>0.22</td>
<td>0.43</td>
<td>0.009</td>
<td>0.008</td>
</tr>
</tbody>
</table>

Notes:
Calculations use uncertainty effect from NLLS regression in column 1 of Table 4. The total effect requires no additional parametric assumptions. Entry growth computed by rescaling the partial effect of export growth according to theoretical model and adding the price index effect. The counterfactual growth from risk reduction evaluates the change in the uncertainty effect, ln(U) and the price index effect, at the long run expected tariff (see text for exact formula).
Table 8: Parameter Values for Quantification and Counterfactuals

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Value</th>
<th>Definition/Application</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\mu$</td>
<td>0.86</td>
<td>Share of manufacturing in total consumption expenditure for consumer welfare quantification</td>
</tr>
<tr>
<td>$I_{t}$</td>
<td>0.045</td>
<td>Import penetration in 2005 to compute price effects, range is [0.022, 0.067] from 2000-10</td>
</tr>
<tr>
<td>$\Delta W_{\tau}$</td>
<td>0.310</td>
<td>Weighted aggregate tariff increase $\sum_V r_V \ln(\tau_{2V}/\tau_{1V})$ for transition from state 1 to 2 use for bounding price change $\ln(P_{2D}/P_{1D})$. Weights are $r_v = \frac{\tau_{1V} R_v}{\sum_V \tau_{1V} R_v}$</td>
</tr>
<tr>
<td>Weighted $\omega$</td>
<td>0.414</td>
<td>Weighted average of profit change $\omega$ over all industries $\sum_V r_V \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-\sigma}$ to compute $d\ln P$</td>
</tr>
<tr>
<td>$\sigma$</td>
<td>3</td>
<td>Median elasticity from Broda and Weinstein (2006), rounded.</td>
</tr>
<tr>
<td>$k$</td>
<td>4.4</td>
<td>Pareto shape parameter, Table 4, NLLS estimates</td>
</tr>
<tr>
<td>$u = \frac{\beta \lambda_{12}}{1 - \beta \lambda_{22}}$</td>
<td>0.636</td>
<td>Expected spell at $s = 2$ for exporter at $s = 1$, NLLS estimate</td>
</tr>
<tr>
<td>$\bar{g}$</td>
<td>1.08</td>
<td>Upper bound price effect adjustment to expected profits in worst case $\left(\frac{P_{2D}}{P_{1D}}\right)^{\sigma - 1}$ for 2005, adjusted by year for counterfactual import penetration 2002-2010</td>
</tr>
<tr>
<td>$\frac{d \ln P}{d \ln \tau}$</td>
<td>0.12</td>
<td>Total elasticity of $P$ to tariffs for welfare equivalents</td>
</tr>
<tr>
<td>$\frac{d \ln P}{d \ln D}$</td>
<td>0.09</td>
<td>Total elasticity of $P$ to transport costs for welfare equivalents</td>
</tr>
</tbody>
</table>

Notes: See data appendix for data sources and online appendix C for exact definitions of expressions used in comparative statics and approximations.
C Online Appendix (not intended for publication)

This Appendix contains a number of results used in the quantification and some intermediate derivations that are useful in proving or deriving other results in the paper. It also contains a notation guide and any tables and figures labeled with the prefix “A”.

C.1 Comparative statics

C.1.1 Effect of trade costs on firm entry and price index

We derive the impact of tariffs and transport costs on the price index and firm entry. These are used in the quantification section to provide an upper bound for the term \( g \) and to calculate the price effect of tariffs and transport costs to include in the GE advalorem equivalent of uncertainty calculation. Recall that

\[ g \equiv (1 - \beta \lambda_{22}) \sum_{T=0}^{\infty} \left( \frac{P_D}{P_T} \right)^{\sigma - 1} \leq \left( \frac{P_D}{P_T} \right)^{\sigma - 1} \text{so } g \leq \exp \left( (\sigma - 1) \ln \frac{P_D}{P_T} \right). \]

We provide a linear approximation to the growth in the deterministic price index due to the change in tariffs from MFN to column 2. We do so in the absence of upgrading and then argue that the expression is similar with upgrading (when written as a function of export shares, which will reflect any upgrading). We show the case for upgrading in section C.6.

Partial elasticity of price index to industry cutoffs and trade costs

Using the full multi-industry price index equation (42) we obtain

\[ \frac{\partial \ln P(c_D(t), \tau)}{\partial \ln P_D} = \frac{1}{1 - \sigma} N_V \int_0^{c_V} \left( \frac{w d_V / \rho}{\rho} \right) \left( \frac{\tau}{\gamma} \right)^{1 - \sigma} \frac{dG(c)}{c_V} = \frac{\tau V R_V}{E} \]

where the last line follows after using \( a = E(1 - \rho)(wd/P\rho)^{1-\sigma} \tau V^{-\sigma} \) and \( R_V = a V \sigma N_V \int_0^{c_V} (c)^{1-\sigma} dG(c) \). The partial tariff elasticity is

\[ \frac{\partial \ln P(c_D, \tau)}{\partial \ln \tau} = \frac{1}{P_{m-\sigma}} N_V \int_0^{c_V} \left( \frac{w d_V / \rho}{\rho} \right) \left( \frac{\tau}{\gamma} \right)^{1 - \sigma} dG(c) = \frac{\tau V R_V}{E} \]

where the last line follows after using \( a = E(1 - \rho)(wd/P\rho)^{1-\sigma} \tau V^{-\sigma} \) and \( R_V = a V \sigma N_V \int_0^{c_V} (c)^{1-\sigma} dG(c) \). Substituting (45) and (46) into the first line of (44), we obtain the total derivative of price index from Appendix A.4. We describe how to use this to quantify changes in the price index using data and estimated parameter below.

Weighted effect of tariff on cutoff

\[ d \ln c_V(c_D, \tau_V) = \frac{\partial \ln c_V(P_D, \tau_V)}{\partial \ln \tau_V} d \ln \tau_V + \frac{\partial \ln c_V(P_D, \tau_V)}{\partial \ln P_D} d \ln P \]

\[ I \sum V \frac{\tau_V R_V}{\Sigma V \tau_V R_V} d \ln c_V(c_D, \tau_V) = \frac{-\sigma}{\sigma - 1} I \sum V \frac{\tau_V R_V}{\Sigma V \tau_V R_V} d \ln \tau_V + I d \ln P \]

where \( I \equiv \frac{\Sigma V \tau_V R_V}{E} \) is Chinese import penetration in total expenditure (see Table 8 for values).

Impact of tariffs on \( P \)
Replacing the weighted effect of tariffs on the cutoff (44) in the equation for the total change in the price index (44), and simplifying, we obtain

\[
d \ln P (c^D, \tau) = \left[ \left( \frac{-\sigma}{\sigma - 1} \right) \frac{k - \sigma + 1}{(1 - \sigma)} + 1 \right] \frac{I}{1 - I \frac{k - \sigma + 1}{(1 - \sigma)}} \sum_V \frac{\tau_m V R_m V}{\sum_V \tau_m V R_m V} \ d \ln \tau_V
\]

If we evaluated the change starting at the MFN values and increasing to column 2 we obtain

\[
d \ln P (c^D (\tau), \tau) |_{\tau_1} = \left[ \frac{\sigma}{\sigma - 1} \frac{(k - \sigma + 1) + (\sigma - 1)}{(k - \sigma + 1) I + \sigma - 1} \right] I \sum_V \frac{\tau_1 V R_1 V}{\sum_V \tau_1 V R_1 V} \ ln \frac{\tau_2 V}{\tau_1 V}
\]

Computing upper bound \( \tilde{g} = \exp \left( (\sigma - 1) \ln \frac{P^D}{P^D} \right) \)

Denoting the weighted tariff change by \( \Delta^W \tau = \sum_V \frac{\tau_1 V R_1 V}{\sum V \tau_1 V R_1 V} \ ln \frac{\tau_2 V}{\tau_1 V} \), we obtain an approximate change in the deterministic price index of

\[
\ln P^D_2 - \ln P^D_1 \approx \left[ \frac{-\sigma}{\sigma - 1} \frac{(k - \sigma + 1) + (1 - \sigma)}{(1 - \sigma) - (k - \sigma + 1) I} \right] \Delta^W \tau \approx 0.037.
\]

We use the price index change approximation to compute the upper bound of \( \tilde{g} = \exp \left( (\sigma - 1) \ln \frac{P^D}{P^D} \right) \). 

Impact of transport cost on \( P \)

This can be similarly found if we note that \( \frac{\partial \ln P (c^D, \tau)}{\partial \ln \tau_V} = \frac{\partial \ln P (c^D, \tau)}{\partial \ln d V} \) and \( d \ln c_V (P^D, d V) |_{\tau} = -d \ln d V + d \ln P \), so

\[
d \ln P (c^D, \tau) \left[ 1 - I \frac{k - \sigma + 1}{(1 - \sigma)} \right] = \left( I \sum_V \frac{\tau V R V}{\sum V \tau V R V} d \ln d V \right) \frac{k}{\sigma - 1} \sum_V \frac{\tau V R V}{\sum V \tau V R V} d \ln d V
\]

Weighted effect of tariff on cutoff

\[
d \ln c_V (P^D, \tau) = \frac{\partial \ln c_V (P^D, \tau_V)}{\partial \ln \tau_V} \ d \ln \tau_V + \frac{\partial \ln c_V (P^D, \tau_V)}{\partial \ln P^D} \ d \ln P^D
\]

\[
I \sum_V \frac{\tau V R V}{\sum V \tau V R V} d \ln c_V (P^D, \tau_V) = \frac{-\sigma}{\sigma - 1} I \sum_V \frac{\tau V R V}{\sum V \tau V R V} d \ln \tau_V + I d \ln P^D
\]

Impact of tariffs on \( P \)

Replacing the cutoff effect above in the price expression and simplifying

\[
d \ln P (c^D, \tau) = \left[ \left( \frac{-\sigma}{\sigma - 1} \right) \frac{k - \sigma + 1}{(1 - \sigma)} + 1 \right] \frac{I}{1 - I \frac{k - \sigma + 1}{(1 - \sigma)}} \sum_V \frac{\tau_m V R_m V}{\sum V \tau_m V R_m V} \ d \ln \tau_V
\]

\[
= \left[ \left( \frac{-\sigma}{\sigma - 1} \right) \frac{k - \sigma + 1}{(1 - \sigma)} + 1 \right] \frac{I}{1 - I \frac{k - \sigma + 1}{(1 - \sigma)}} \sum_V \frac{\tau_m V R_m V}{\sum V \tau_m V R_m V} \ d \ln \tau_V
\]
If we evaluated the change starting at the MFN values and increasing to column 2 we obtain
\[ d \ln P \left( e^P (\gamma), \tau_1 \right) |_{\tau_1} = \frac{\sigma \gamma}{\sigma - 1} \frac{(k - \sigma + 1) + \sigma - 1}{k - \sigma + 1} \sum \frac{\tau_1 V R_{1V}}{\tau_1 V} \ln \frac{\tau_2 V}{\tau_1 V} \]

C.1.2 Effect of policy uncertainty on firm entry and price index

In equation (18) we provide the semi-elasticity of the price index wrt γ as a function of \( \varepsilon_V \equiv \partial \ln P (c) / \partial \ln c^V \). We now show that this is similar to the deterministic elasticity derived in section C.1.1 and then use it to obtain the relationship between \( P (\cdot) \) and γ in terms of model parameters and data. We show the expressions are similar in the presence of upgrading in section C.6. We also use the expression to evaluate the GE impact of changes in uncertainty on entry.

Using the equilibrium price paid by consumers of imported goods, \( p_o = (w V d V c_o / \rho) \), and the Pareto distribution in (21) we then obtain \( \varepsilon_V \) with TPU as follows

\[
\frac{\partial \ln P (c)}{\partial \ln c^V} = (1 - \sigma)^{-1} \tau_N \left[ k \left( d V w / (P \rho) \right)^{\sigma} \frac{N_V}{c^V} (c^V)^{\kappa - \sigma + 1} \right]
\]

\[ \equiv - k - \sigma + 1 \frac{\tau_N R_V}{\sigma - 1} \frac{E}{\varepsilon_V} = \varepsilon_V \quad (48) \]

The last line is obtained by rearranging the equilibrium expression for exports to replace the term in brackets. This compact expressions shows the price index elasticity to an industry cutoff is proportional to its share of expenditure.

Semi-elasticity of P wrt γ

Using the expression for \( \varepsilon_V \) derived above in equation (18) and noting that \( \tilde{\varepsilon}_V \equiv \frac{\varepsilon_V}{(1 - \sum \varepsilon_V)} \) we obtain

\[
\left. \frac{d \ln P_1 (c_1)}{d \gamma} \right|_{\gamma = 0} = \frac{\beta \lambda_2}{(\sigma - 1) (1 - \beta \lambda_{22})} \sum \tilde{\varepsilon}_1 V (\tilde{\omega} - 1)
\]

\[ = \frac{\beta \lambda_2}{(\sigma - 1) (1 - \beta \lambda_{22})} \sum \frac{k - \sigma + 1 \tau_1 V R_{1V}}{1 - \frac{k - \sigma + 1}{1 + \sigma} \sum \tau_1 V R_{1V}} (\tilde{\omega}_1 V - 1)
\]

\[ = \frac{\beta \lambda_2}{(\sigma - 1) (1 - \beta \lambda_{22})} \frac{(k - \sigma + 1) I_1}{(k - \sigma + 1) I_1 + \sigma - 1} \sum \tau_1 V R_{1V} (1 - \tilde{\omega}_1 V)
\]

where \( I_1 \equiv \frac{\sum \tau_1 V R_{1V}}{E} \) and \( r_{1V} \equiv \frac{\tau_1 V R_{1V}}{\sum \tau_1 V R_{1V}} \). Thus a necessary and sufficient condition for \( \left. \frac{d \ln P_1 (c_1)}{d \gamma} \right|_{\gamma = 0} > 0 \) is \( \sum \tau_1 V (1 - \tilde{\omega}_1 V) |_{\gamma = 0} > 0 \).

Semi-elasticity of entry wrt γ

\[
\sum \tau_1 V \frac{d \ln c^V}{d \gamma} \bigg|_{\gamma = 0} = \sum \tau_1 V \left[ \left. \frac{d \ln U_1 (\tilde{\omega})}{d \gamma} \right|_{\gamma = 0} + \left. \frac{d \ln P_1 (c_1)}{d \gamma} \right|_{\gamma = 0} \right]
\]

\[ = \frac{\beta \lambda_2}{(\sigma - 1) (1 - \beta \lambda_{22})} \left[ \sum \tau_1 V (\tilde{\omega} - 1) + \frac{(k - \sigma + 1) I_1}{(k - \sigma + 1) I_1 + \sigma - 1} \sum \tau_1 V (1 - \tilde{\omega}) \right]
\]

\[ = \frac{\beta \lambda_2}{(\sigma - 1) (1 - \beta \lambda_{22})} \left[ 1 - \frac{(k - \sigma + 1) I_1}{(k - \sigma + 1) I_1 + \sigma - 1} \right] \sum \tau_1 V (\tilde{\omega} - 1)
\]
C.2 Ordering of Cutoffs

\[ c_2^U = c_2^D < c_m^U < c_m^D \leq c_0^U = c_0^D \]

In the proof of proposition 1 we provide the conditions for \( c_1^U < c_1^D \leq c_0^U = c_0^D \). In the text we showed that \( c_2^U = c_2^D \). Thus here we prove the statement in the text that \( c_2^U < c_1^U \) if and only if \( \omega < 1 \).

\[
\begin{align*}
  c_2^U &= c_2^D = \left[ \frac{a_2}{(1 - \beta) K} \right]^{\frac{1}{\sigma'}} < c_1^D \left( \frac{1 + u(\gamma) \times \left( \frac{\tau_2}{\tau_1} \right)^{-\sigma}}{1 + u(\gamma)} \right)^{\frac{1}{\sigma'}} \\
  \left[ \frac{a_2}{(1 - \beta) K} \right]^{\frac{1}{\sigma'}} &< \left[ \frac{a_1}{(1 - \beta) K} \right]^{\frac{1}{\sigma'}} \left( \frac{1 + u(\gamma) \times \left( \frac{\tau_2}{\tau_1} \right)^{-\sigma}}{1 + u(\gamma)} \right)^{\frac{1}{\sigma'}} \\
  \left( \frac{\tau_2}{\tau_1} \right)^{-\sigma} &< \left( \frac{1 + u(\gamma) \times \left( \frac{\tau_2}{\tau_1} \right)^{-\sigma}}{1 + u(\gamma)} \right)
\end{align*}
\]

This can simplified to yield the inequality \( u(\gamma)\omega < u(\gamma) \) that we showed holds iff \( \tau_2 > \tau_1 \) and \( \omega < 1 \) in the proof of Proposition 1.

C.3 Price index expectations and transition dynamics

Expectations of future price index \( P_s^e \)

Firms can derive \( P_s^e \) as follows. First, they can correctly infer the last term in the RHS of \( 9 \)—the domestic component of \( P_s \)—because (i) in each country there is a time invariant distribution from which firms can be born and (ii) there is a constant mass of potential blueprints \( (c_v) \) so when a fraction dies a similar fraction is born and (iii) there are no domestic fixed costs of production so all potential varieties are produced and sold domestically. Second, to predict the first term of \( P_s \) firms use the observed policy realization, \( \tau_m \), and must infer the set of exported varieties, \( \Omega_x^s \), over which to integrate. The latter is simply \( \Omega_x^s = \Omega_x^{cont} \cup \Omega_x^{entry} \) where \( \Omega_x^{cont} \) represents the set of foreign producers that exported to this market both in the previous and current periods (so \( \Omega_x^{cont} = \emptyset \) in the initial trading period). The measure of continuers is given by the measure of previous period exporters—observed in \( \Omega_{t-1} \)—adjusted by the exogenous survival probability, \( \beta \), applied to all subsets. So \( \Omega_x^{cont} \) is independent of the current tariff and economic conditions. New exporters are represented by the subset \( \Omega_x^{entry} \) of all potential firms in the foreign country that (i) did not export in the previous period—known from \( \Omega_{t-1} \)—and (ii) have a cost such that entry is optimal in state \( s \) according to \( 4 \).

Transition Dynamics

The price index as a function of \( T \) is

\[
P_{2T}^D = \left[ N \tau_2 \left( \int_0^{\tau_2} (c_v/\rho)^{1-\sigma} dG(c) + \beta^T \int_{\tau_2}^{\tau_1} (c_v/\rho)^{1-\sigma} dG(c) + \int_{\Omega_x^h} (c_v/\rho)^{1-\sigma} \, dv \right) \right]^{1/(1-\sigma)} \tag{49}
\]

In the text we claim \( P_{2T}^D \leq P_2^D \) for all \( T \). To verify this (and derive an expression for the transition dynamics...
for the price index when  amalgamate (11) in (49), totally differentiating and simplifying we have
\[
\frac{d \ln P^D_{2T}}{dT} = \frac{d \ln P^D_{2T}}{dT} = \left(\frac{P^D_{2T}}{\rho} \right)^{\sigma - 1} \left[ N\tau_2 \left(1 - \beta^T \right) (c^D_{2T}/\rho)^{1-\sigma} g(c^D_{2T}) \frac{dc^D_{2T}}{dT} + \left( \int c^D_{2T} (c_v/\rho)^{1-\sigma} dG(c) \right) \beta^T \ln \beta \right]
\]
\[
= \left(\frac{P^D_{2T}}{\rho} \right)^{\sigma - 1} \left[ N\tau_2 \left(1 - \beta^T \right) (c^D_{2T}/\rho)^{1-\sigma} g(c^D_{2T}) \frac{dc^D_{2T}}{dT} + \left( \int c^D_{2T} (c_v/\rho)^{1-\sigma} dG(c) \right) \beta^T \ln \beta \right]
\]
\[
= \left(\frac{P^D_{2T}}{\rho} \right)^{\sigma} N\tau_2 (\beta^T) \left[ \int c^D_{2T} (c_v/\rho)^{1-\sigma} dG(c) \right] + (1 - \sigma)(1 - \eta_t) > 0
\]

The first line uses the fact that the mass of domestic firms, 0, is time invariant and applies the Leibniz rule to the remaining terms. The second line: \( dc^D_{2T}/dT \) because at a fixed tariff \( c^D_{2T} \) only changes over time due to price index changes. The third line re-arranges using \( 1 - \eta_t \equiv 1 - N\tau_2 (P^D_{2T})^{-1} (1 - \beta^T) (c^D_{2T}/\rho)^{1-\sigma} g(c^D_{2T}) c^D_{2T} > 0 \). The inequality follows from \( \ln \beta < 0 \) (so negative numerator) and \( (1 - \sigma)(1 - \eta_t) < 0 \) because \( 1 - \sigma < 0 \) and \( 1 - \eta_t > 0 \).

C.4 Additional Welfare Results

Quantification of expected welfare change from MFN to agreement state

The quantification focuses on the impact of uncertainty on the growth of welfare in the MFN state, i.e. \( -\mu \gamma \frac{d \ln P_1}{d\gamma} \bigg|_{\gamma=0} \). We can also calculate the effect of uncertainty on expected welfare growth based on our parameters.

\[
\ln \frac{W(\tau = \tau_0, \gamma > 0)}{W(\tau = \tau_0, \gamma = 0)} \bigg|_{\tau_0 \to \tau_1} = \ln \frac{W(\gamma > 0)}{W(\gamma = 0)}
\]
\[
= -\mu \gamma \frac{d \ln P_1}{d\gamma} \bigg|_{\gamma=0} + \beta \lambda_{12} \frac{\beta}{1 - \beta \lambda_{22}} \times \ln \mathbb{E}_{s \in \mathbb{E}} \left( \frac{P^D_s}{P^D_1} \right)^{\mu}
\]
\[
= -\mu \gamma \frac{d \ln P_1}{d\gamma} \bigg|_{\gamma=0} + u \times \ln \mathbb{E}_{s \in \mathbb{E}} \left( \frac{P^D_s}{P^D_1} \right)^{\mu}
\]

The first line uses the fact that \( \tau_1 \sim \tau_0 \) in this setting. The second uses the approximation in (20) and \( \ln \mathbb{E}_{s \in \mathbb{E}} \left( \frac{P^D_s}{P^D_1} \right)^{\mu} \bigg|_{\tau_0 \to \tau_1} = 0 \). The third uses the definition of \( u \). The first term is the with state effect reported in the text. The second term is obtained using our estimate of \( u \) and \( \ln \mathbb{E}_{s \in \mathbb{E}} \left( \frac{P^D_s}{P^D_1} \right)^{\mu} \) (see Appendix C.1.1), which is a lower bound for \( \ln \mathbb{E}_{s \in \mathbb{E}} \left( \frac{P^D_s}{P^D_1} \right)^{\mu} \).

Welfare Certainty Equivalents

We define a percent change in trade costs, \( \Delta \), as the trade cost equivalent of TPU if it yields the same change in welfare as eliminating uncertainty at given tariffs. So for tariffs, the value of \( \Delta \) computed in section 3.7 is the implicit solution to

\[
W(\tau = \tau_1 (1 + \Delta), \gamma = 0) = W(\tau = \tau_1, \gamma > 0)
\]
We solve for $\Delta$ using the following first order log approximations
\[
\ln W(\tau = \tau_1 (1 + \Delta), \gamma = 0) \approx \ln W(\tau = \tau_1, \gamma = 0) + \Delta \frac{\partial \ln W(\tau, \gamma = 0)}{\partial \ln \tau}|_{\tau_1} \\
\ln W(\tau = \tau_1, \gamma > 0) \approx \ln W(\tau = \tau_1, \gamma = 0) + \gamma \frac{\partial \ln W(\tau = \tau_1, \gamma > 0)}{\partial \gamma}|_{\gamma=0}
\]
Then by equating the two equations above and solving for $\Delta$ we obtain of a quantifiable expression for the certainty equivalent
\[
\Delta \approx \gamma \times \left. \frac{\partial \ln W(\tau = \tau_1, \gamma > 0)}{\partial \gamma} \right|_{\gamma=0} / \left. \frac{\partial \ln W(\tau, \gamma = 0)}{\partial \ln \tau} \right|_{\tau_1}
\]

**Risk Decomposition of Welfare Changes**

In footnote 67 we state that $-\mu \gamma \frac{d\ln P}{d\gamma}|_{\gamma=0}$ evaluated at the long-run tariffs provides an upper bound for the first order risk effect of TPU on expected welfare. The effect of TPU on expected welfare is given by (20) so here we establish that the second term on the RHS is non-negative to a first order approximation when evaluated at $\bar{\tau}$.

\[
\sum_{m=0,2} \beta \lambda_{1m} \ln \mathbb{E}_{s \in m} \left( \frac{P^D_m}{P^D_1} \right)^{-\mu} > \sum_{m=0,2} \beta \lambda_{1m} \ln \left( \frac{P^D_m}{P^D_1} \right)^{-\mu} \\
\sum_{m=0,2} \beta \lambda_{1m} \ln \left( \frac{P^D_m}{P^D_1} \right)^{-\mu} \biggr|_{\lambda_{mm} \to 1, \ln \tau_1 = \Sigma \lambda_m \ln \tau_m} = -\mu \sum_{m=0,2} \lambda_m \ln \left( \frac{P^D_m}{P^D_1} \right)^{-\mu} \biggr|_{\ln \tau_1 = \Sigma \lambda_m \ln \tau_m} \approx 0
\]
The inequality in the first line follows by replacing the higher steady state price index for each of the transition values under column 2, $P^D_m$, with an absorbing state then the second term on the RHS is non-negative to a first order when evaluated at $\bar{\tau}$.

The equality in the second line evaluates at $\lambda_{mm} \to 1$, as we do when computing $\bar{\tau}$, and uses $\lambda_{1m} \equiv \gamma \lambda_m$ and $\Sigma \lambda_m = 1$ for $m = 0, 2$.

The last expression is zero to a first order when we approximate $P^D_m$ in $m = 0, 2$ around $\tau_1$ and use $\Sigma \lambda_m = 1$.

The calculation in the text evaluates $-\mu \gamma \frac{d\ln P}{d\gamma}|_{\gamma=0}$ at $\bar{\tau} = \Sigma \lambda_m \tau_m$ but the result is very similar if we instead used $\ln \tau_1 = \Sigma \lambda_m \ln \tau_m$ because $\tau - 1^{-\ln \tau}$ for most industries

### C.5 Expected value as duration weighted average of payoffs

In the proof of proposition 3 we interpret the weights $n_m$ in (19) as the expected duration in $m$ conditional on starting at $m = 1$. Here we prove this interpretation.

Writing the system of equations in (40) in matrix form, $(I - \beta M) W = w$, we have
\[
\begin{bmatrix}
1 - \beta \lambda_{22} & -\beta \lambda_{21} & -\beta \lambda_{20} \\
-\beta \lambda_{12} & 1 - \beta \lambda_{11} & -\beta \lambda_{10} \\
-\beta \lambda_{02} & -\beta \lambda_{01} & 1 - \beta \lambda_{00}
\end{bmatrix}
\begin{bmatrix}
W_2 \\
W_1 \\
W_0
\end{bmatrix} =
\begin{bmatrix}
w_{12} \\
w_{11} \\
w_{10}
\end{bmatrix} = W = Nw
\]
where $M$ denotes the $3 \times 3$ policy matrix with typical term $\lambda_{mm'}$ from (3), $W$ and $w$ the $3 \times 1$ vectors associated with $W_m$ and $w_m$. By solving for the second row of $N \equiv (I - \beta M)^{-1}$ we obtain the weights in (19). To interpret those weights as expected number of periods we make use of the following result:

If $N$ is the fundamental matrix for a Markov process with matrix $P$ with an absorbing state then the typical entry $n_{ij}$ represents the expected time that the process is in transient state $j$ conditional on starting
at \( i \) where the initial state is counted if \( i = j \) Consider the relevant Markov process for individuals that survive with probability \( \beta \in (0, 1) \). It is characterized by the transition matrix \( P \), which augments the policy matrix with the death state.

\[
P = \begin{bmatrix}
\beta \lambda_{22} & \beta \lambda_{21} & \beta \lambda_{20} & 1 - \beta \\
\beta \lambda_{12} & \beta \lambda_{11} & \beta \lambda_{10} & 1 - \beta \\
\beta \lambda_{02} & \beta \lambda_{01} & \beta \lambda_{00} & 1 - \beta \\
0 & 0 & 0 & 1
\end{bmatrix} = \begin{bmatrix}
\beta M & I - \beta \\
0 & 1
\end{bmatrix}
\]

Using standard results we have that the probability of being in a given state after \( t \) periods is \( P^t = [((\beta M)^t) \ast \mathbf{1}] \) where \( \ast \) denotes a \( 3 \times 1 \) vector that we do not require. Moreover, that probability is goes to zero for any policy state because they are transient due to the positive probability of death, i.e. \( 1 - \beta > 0 \Rightarrow \lim_{t \to \infty} (\beta M)^t = 0 \). The probability of given state \( m' \) after starting at \( m \) after \( t \) periods is given by the element of \( P^t \) denoted by \( (\beta M_{mm'}) \) so the expected number of periods in \( m' \) after starting at \( m \) is equal to \( \Sigma_{t=0,\infty} (\beta M)^t \). Therefore, since the inverse matrix \( N \equiv (I-\beta M)^{-1} = \Sigma_{0,\infty} (\beta M)^t \), and so \( n_{ij} \) is the expected number of times that the process is in transient state \( j \) conditional on starting at \( i \).

C.6 Effect of policy uncertainty on upgrading and price index

Similarly to the derivation without upgrading we split the price index into the subcomponents depending on the country of origin and industry (1st line below). In the second line we further divide the foreign varieties in to the endogenous set of firms that upgrades \( (\Omega_{V,ch}) \), since they will have lower equilibrium prices and the remaining set of firms \( (\Omega_{V,ch} \setminus \Omega_{V,ch}) \).

\[
P^{1-\sigma} = \int_{v \in \Omega} (p_v)^{1-\sigma} dv = \sum_{V,ch} \int_{v \in \Omega_{V,ch}} (p_v)^{1-\sigma} dv + \sum_{V,C \neq ch} \int_{v \in \Omega_{V,C}} (p_v)^{1-\sigma} dv
\]

Using the equilibrium price, \( p_v = (w_{TV} d_{TV} c_v / \rho) \) for the non-upgraders and \( z (w_{TV} d_{TV} c_v / \rho) \) for the upgraders, as well as the Pareto distribution we obtain \( \varepsilon_V \) at a state such as the MFN one where the cutoffs are \( c_{mz} = \phi_{c m} \) (we omit the state subscript below and the \( V \) subscripts for the technological parameters for notational simplicity)

\[
\frac{\partial \ln P (c)}{\partial \ln c_V} \bigg|_{c_V} = (1-\sigma)^{-1} \frac{1}{P^{1-\sigma}} \left[ \frac{N_V}{c_k} \left( w_{TV} d_{TV} / \rho \right)^{1-\sigma} k \left( \frac{z^{1-\sigma}}{\phi_{c V}^k} (c_{TV})^{k-\sigma+1} + \left( \frac{c_{TV}}{\rho} \right)^{k-\sigma+1} - \left( \frac{c_{TV}}{\rho} \right)^{k-\sigma+1} \right) \right]
\]

where \( \zeta_V = 1 + (\phi_V)^{1-\sigma+k} (z^{1-\sigma} - 1) \) and the expression in brackets the same we derived for \( \varepsilon_V | \zeta_V = 1 \), i.e. without upgrading so \( \varepsilon_V = \varepsilon_V | \zeta_V = 1 \zeta_V \). This implies that the price elasticity wrt each cutoff is higher under upgrading. Since we also have that exports with upgrading can be written similarly: \( R_V = R_V | \zeta_V = 1 \zeta_V \) we obtain the same general expression for the elasticity when written in terms of the export value

\[
\varepsilon_V \equiv \frac{\partial \ln P (c)}{\partial \ln c_V} \bigg|_{c_V} = \frac{-k-\sigma+1 \tau_V R_V / \rho}{\sigma-1} \frac{\sigma}{E}
\]

**Semi-elasticity of \( P \) wrt \( \gamma \) under upgrading**

\( P \) depends on \( \gamma \) only via \( c_V^U \) so \( d \ln P / d \gamma \) is the same as derived without upgrading but now \( \varepsilon_V \) reflects

---

\(^{70}\)For a full proof see “Introduction to Probability” Theorem 11.4 at <http://www.dartmouth.edu/chance>
any upgrading that took place, as embodied in \( R_V \), that is we still obtain

\[
\frac{d \ln P_1}{d \gamma} \bigg|_{\gamma=1} = \frac{\beta \lambda_2}{(k-\sigma+1) I} \left( \frac{(k-\sigma+1) I}{(k-\sigma+1) I + \sigma - 1} \sum_v r_{1V} \right)
\]

where \( I \equiv \frac{\sum_v r_{1V}}{\sum_v R_{1V}} \) and the tariff inclusive import weights evaluated under the MFN state are defined as \( r_{1V} \equiv \frac{\sum_v R_{1V}}{\sum_v R_{1V}} \).

Semi-elasticity of entry and upgrading wrt \( \gamma \)

We also obtain a similar expression in terms of export revenues as in the absence of upgrading for the weighted semi-elasticity of entry with respect to uncertainty

\[
\sum_v r_{1V} \frac{d \ln e_{1V}^U}{d \gamma} \bigg|_{\gamma=0} = \left( 1 - \frac{(k-\sigma+1) I}{(k-\sigma+1) I + \sigma - 1} \right) \frac{\beta \lambda_2}{(\sigma-1)(1-\beta \lambda_2)} \sum_v r_{1V} (\hat{\omega}_V - 1)
\]

which also applies to the upgrading cutoff since \( c_{1z}^U = \phi c_1^U \).

C.7 EU-15 and Taiwan Falsification Checks

The data for Table A3 come from two sources. Taiwan’s exports and transport cost measures to the U.S. are obtained at the HS-6 level using disaggregated trade data from the NBER. Taiwan is eligible for preferential rates, but over 99% of Taiwan’s exports in all years from 1996-2006 receive MFN tariff treatment. For the E.U.-15, we use aggregated HS-6 imports from China from COMTRADE. MFN tariff data were obtained from TRAINS. All these data are concorded to the 1996 HS revision for consistency over time.

C.8 Double difference specification

If there is an industry specific growth rate trend in the number of firms, \( \Delta (\ln N_V) = \theta_V \), and \( \theta_V \) is correlated with our policy or trade cost variables, then identification is still possible via a difference-of-differences approach. We illustrate this using the mass of firms but the variables in \( n_V \) could also be allowed to have industry specific time trends. This yields the following long difference

\[
\Delta_{10} \ln R_V = b_\gamma \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) + b_\tau \Delta \ln \tau_V + b_d \Delta \ln D_V + b + \theta_V + u_V
\]

where \( \Delta_{10} \) is subscripted to denote the difference over a transition from 1 to 0.

Now consider taking the difference between two years that remain in state 1. The difference above uses 2000 (1) and 2005 (0), but we can also use the difference between 1999(1) and 1996(1) and denote it by \( \Delta_{11} \)

\[
\Delta_{11} \ln R_V = -\Delta_{11} b_\gamma \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) + b_\tau \Delta_{11} \ln \tau_V + b_d \Delta_{11} \ln D_V + b' + \theta_V + u_V \tag{51}
\]

Since both our uncertainty measure and the estimated parameters on the uncertainty measure could change over time, we denote the parameter on uncertainty by \( b_\gamma' \) and note that there are two components to the change in the first term

\[
-\Delta_{11} b_\gamma' \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) = -b_\gamma' \Delta_{11} \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) - \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) \Delta_{11} b_\gamma
\]

The second term is evaluated at final period tariffs, which are very close to 2000 levels. Because \( \tau_{2V} \) is fixed during this period and any variation in \( \left( \frac{\tau_{2V}}{\tau_{1V}} \right) \) is due to small changes in \( \tau_{1V} \), already controlled by \( \Delta_{11} \ln \tau_V \).
, we take $\Delta_{11} \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) \approx 0$ to obtain

$$-\Delta_{11} b'_{\gamma} \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) \approx - \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) \Delta_{11} b'_{\gamma}$$

$$= - \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) k - \sigma + 1 \beta \lambda_{12} \frac{1}{\sigma - 1} \Delta_{11} g'_{1} = - \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) b'_{\gamma} g_{1} \Delta_{11} g'_{1}.$$  

We then take the double difference, normalizing each differenced RHS variable by the length of the time period to obtain magnitudes comparable to our first differenced results

$$\frac{\Delta_{10} \ln R_{V}}{5} - \frac{\Delta_{11} \ln R_{V}}{3} = b_{\gamma} \left( 1 + \frac{\Delta_{11} g'_{1}}{g_{1}} \right) \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) + b_{\tau} \left( \frac{\Delta_{10} \ln \tau_{V}}{5} - \frac{\Delta_{11} \ln \tau_{V}}{3} \right) + b_{d} \left( \frac{\Delta_{10} \ln D_{V}}{5} - \frac{\Delta_{11} \ln D_{V}}{3} \right) + b - b' + u_{V} - u'_{V}$$

(52)

(53)

The coefficients from estimating equation (53) have the same interpretation as our OLS baseline. The sample size drops since we can only use HS6 industries traded in 2005, 2000, 1999, and 1996. Further, the double differenced variables are somewhat noisy so we employ a robust regression routine that downweights outliers more than 7 times the median absolute deviation from the median residuals, iterating until convergence.

### C.9 Yearly panel specification

The full panel specification used to obtain the coefficients in Figure A1 allows us to examine how the uncertainty coefficient changed over time. Consider a generalized version of the level equation (23) that allows the uncertainty coefficient to vary by year, $t$, and includes time by sector effects, $b_{iS}$, in addition to industry (HS-6) fixed effects $b_{V}$.

$$\ln R_{TV} = -b_{\gamma t} \left( 1 - \left( \frac{\tau_{2V}}{\tau_{1V}} \right)^{-\sigma} \right) + b_{\tau} \ln \tau_{TV} + b_{d} \ln D_{TV} + b_{iS} + b_{V} + u_{TV} \ ; \ t = 1996 \ldots 2006$$

We estimate two versions of this equation. First, recall that there is almost no variation over 2000-2005 in the uncertainty variable so in the baseline we focused in the change in coefficient. To compare the panel results with the baseline we initially use $\tau_{tV} = \tau_{2000V}$ to construct the uncertainty measure. In this case we cannot identify $b_{\gamma t}$ for each year since the uncertainty regressor only varies across $V$ and we include $b_{V}$. Instead, we estimate the coefficient change over time relative to a base year, namely $b'_{\gamma t}^\text{panel} = - (b_{\gamma t} - b_{\gamma 2000}) = \frac{k - \sigma + 1}{\sigma - 1} \beta \lambda_{12} \Delta_{\gamma t}$, where $\Delta_{\gamma t} = \gamma_{2000} - \gamma_{t}$. We obtain similar results to Figure A1 if we drop the year 2001, constrain $b'_{\gamma t}^\text{panel}$ to a single value for pre-WTO and a single value post-WTO, or both. All results available upon request.

### C.10 Entry specification and quantification

There is additional evidence for the entry uncertainty channel when we explore more detailed data. We can identify each of the coefficients in (30) up to a factor, $\nu' \in [0, 1]$, by using growth in product counts within each HS-6 industry. Given the data limitations, we focus on linear estimates, approximate the uncertainty term around $\gamma = 0$

The change in the number of exported varieties is unobserved to us and so we treat it as a latent variable and model how it maps to observable changes in exported products. We assume there is a continuous, increasing,
differentiable function \( \nu(\cdot) \) that maps varieties to product counts: \( \ln(pcount_{sV}) = \nu(\ln(number_{sV})) \). If there was only one firm in an HS-6 industry and it produced a single variety then we would observe one traded product (an HS-10 category) within an industry. We cannot observe more traded products than the maximum number tracked by customs in each industry, i.e. the total number of HS-10 categories in an HS-6. So clearly we have a lower bound \( \nu(\ln(number_{sV}) = 0) = 0 \) and an upper bound \( \ln(pcount_{sV}) = \nu(\ln(number_{hV})) \) for all \( \ln(number_{tV}) \) at least as high as \( \ln(number_{hV}) \)—the (unobserved) threshold where all HS-10 product categories in an HS-6 industry have positive values. If we assume product counts and varieties are continuous, then \( \nu' \geq 0 \) for \( number_{V} \in (0, number_{hV}) \) and zero otherwise. The weak inequality accounts for the possibility that different firms export within the same HS-10 category so there is true increase in variety that is not reflected in new HS-10 categories traded. We log linearize the equation of product counts around \( \ln(number_{t-1V}) \). Then the change in products between \( t \) and \( t - 1 \) is \( \Delta \ln(pcount_{V}) = \nu'(\ln(number_{t-1V}) \Delta \ln number_{V} + b^e \Delta \ln D_{V} + b^e + e_{V} \) (54)

The estimation coefficients obtained from a linear regression rescale the parameters in (30) by \( \nu' \). The predicted coefficients are \( b^e = \frac{k}{\sigma-1} u g u \nu' \geq 0, b^e = \frac{\sigma k}{\sigma-1} \nu' \leq 0 \) and \( b^d = -k \nu' \leq 0 \) and the constant \( b^e = -b^e (1-g^{-1}) + \nu' \frac{k}{\sigma-1} \Delta \ln A \). The weak inequalities capture the possibility that \( \nu' = 0 \).

In going from this specification to the data, we account for the maximum number of tradable products within each HS-6. We use this information to impose sample restrictions on the regression consistent with our specification of the \( \nu(\cdot) \) function. We drop observations if an industry already trades the maximum number of products available at the beginning and end of the sample—a “max-to-max” transition where \( \nu' = 0 \) as well as industries that are non-traded throughout the sample—“zero-to-zero” transitions. This means we have a symmetric sample restriction at the upper and lower bounds suggested by our mapping \( \nu \).

In estimating the log changes specification we must focus only on the industries with some traded product in both periods, which is what we also did in the baseline trade flow regression. In Table A5, we report the results of the specifications analogous to Table 2 but focusing on variety growth. The first column shows the baseline specification and we find that all three variables have the predicted sign and are statistically significant. In column 3 we control for sector effects and find similar results. In both cases we test and fail to reject that \( b^e / b^d = \sigma / (\sigma - 1) \) and impose this constraint in columns 2 and 4. The point estimates of the tariff and transport cost elasticity are lower than their respective values in the export equation (Table 2 column 4), which is consistent with the fact that the parameters here are scaled by \( \nu' \in [0,1] \). In the working paper we also compare these estimates for entry with their linear counterpart for exports and find that TPU has a higher impact on entry than exports, which is consistent with the model’s predictions.
## Table A1: Summary Statistics Across Regression Specifications

<table>
<thead>
<tr>
<th></th>
<th>1-3</th>
<th>4</th>
<th>A5</th>
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</thead>
<tbody>
<tr>
<td>Chinese export growth to US ($Δ\ln$, 2005-2000)</td>
<td>1.29</td>
<td>1.27</td>
<td>n/a</td>
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<tr>
<td></td>
<td>[1.672]</td>
<td>[1.604]</td>
<td></td>
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<td>Uncertainty Pre-WTO (2000)</td>
<td>0.52</td>
<td>0.52</td>
<td>0.52</td>
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<tr>
<td></td>
<td>[0.202]</td>
<td>[0.198]</td>
<td>[0.193]</td>
</tr>
<tr>
<td>Change in Tariff ($Δ\ln$)</td>
<td>-0.003</td>
<td>-0.003</td>
<td>-0.005</td>
</tr>
<tr>
<td></td>
<td>[0.00884]</td>
<td>[0.009]</td>
<td>[0.0114]</td>
</tr>
<tr>
<td>Change in Transport Costs ($Δ\ln$)</td>
<td>-0.005</td>
<td>-0.001</td>
<td>-0.007</td>
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<td></td>
<td>[0.0870]</td>
<td>[0.041]</td>
<td>[0.0861]</td>
</tr>
<tr>
<td>Change in MFA quota status (binary)</td>
<td>-0.129</td>
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<td></td>
<td>[0.335]</td>
<td></td>
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<tr>
<td>Change in TTB status (binary)</td>
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<td>n/a</td>
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<td>Ratio of Column 2 to MFN Tariff</td>
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<td>Product growth</td>
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<td>n/a</td>
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<tr>
<td></td>
<td></td>
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<td>[0.463]</td>
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<p>| | | | |</p>
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<tr>
<th></th>
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</thead>
<tbody>
<tr>
<td>Observations</td>
<td>3,242</td>
<td>3,074</td>
<td>1,227</td>
</tr>
<tr>
<td>Fraction of total export growth</td>
<td>0.976</td>
<td>0.974</td>
<td>0.262</td>
</tr>
</tbody>
</table>

**Notes:**
- Means with standard deviation in brackets. See referenced table and text for detailed information about sample and variable definitions. "n/a": not applicable since variable not used in the corresponding table.
Table A2: Export Growth from China (2000-2005) Robustness

<table>
<thead>
<tr>
<th>Potential Issue</th>
<th>Magnitude of common σ</th>
<th>Industry variation in σ</th>
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</thead>
<tbody>
<tr>
<td>Estimation</td>
<td>OLS, σ=3</td>
<td>OLS, σ=2</td>
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<tr>
<td>Sample change vs. baseline</td>
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<td>none</td>
</tr>
<tr>
<td>Observations</td>
<td>3242</td>
<td>3242</td>
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<tr>
<td>R-squared</td>
<td>0.028</td>
<td>0.027</td>
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<tr>
<td>Sector fixed effects</td>
<td>no</td>
<td>yes</td>
</tr>
<tr>
<td>Restriction p-value (F-test)</td>
<td>0.195</td>
<td>0.311</td>
</tr>
</tbody>
</table>

Notes: Robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.10. Predicted sign of coefficient in brackets under variable.

(…) Constant or sector fixed effects included as noted. Tariff and transport cost changes included but not reported for space considerations. The typical coefficient is b_t= -2.5 for transport cost and b_t= b_d(σ/(σ-1)) can’t be rejected at p-values listed in last row. Uncertainty similar to Table 2 with σ=3 except in Panel A (uses listed values, σV = median estimate within HS6) and Panel B columns 5 and 6.
Panel B columns 5 and 6 : use both ad-valorem tariff and the advalorem equivalent of specific tariffs (AVE=specific tariff / unit value).
Panel B columns 1 and 2: Robust regression downweights outliers more than 7 times the median absolute deviation from the median residual.
Panel B: columns 3 and 4: Midpoint growth of export level R is given by 2*(R(t)-R(t-1))/(R(t)+R(t-1)) for t=2005 and t-1=2000. Defined for all exported 6 digit HS codes Panel B: columns 7 and 8 drop HS Section XVI: machinery and electrical applications; electrical equipment; parts thereof; sound recorders and reproducers, television image and sound recorders and reproducers, and ports and accessories of such articles.
Table A3: Falsification test for unobserved HS-6 level shocks to Chinese export supply and US import demand (2000-2005)

<table>
<thead>
<tr>
<th>Dependent variable</th>
<th>Chinese export growth to:</th>
<th>U.S. import growth from:</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>USA</td>
<td>EU-15</td>
</tr>
<tr>
<td>1</td>
<td></td>
<td></td>
</tr>
<tr>
<td>2</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Uncertainty pre-WTO (US)</td>
<td>0.461** [0.192]</td>
<td>0.0244 [0.184]</td>
</tr>
<tr>
<td>Change in bilateral transport cost</td>
<td>n/a</td>
<td>-3.744*** [0.665]</td>
</tr>
<tr>
<td>Observations</td>
<td>3,041</td>
<td>3,041</td>
</tr>
<tr>
<td>R-squared</td>
<td>0.031</td>
<td>0.038</td>
</tr>
<tr>
<td>Sector Fixed Effects</td>
<td>yes</td>
<td>yes</td>
</tr>
</tbody>
</table>

Notes:
- Robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. Uncertainty pre-WTO is defined as in the baseline US sample. Change in importer tariff in column 2 is the EU-15 MFN tariff change and in other columns it is the US MFN tariff change (similar for Taiwan and China).
- Transport cost data for EU is unavailable. For comparability we also omit that variable for the US regression in column 2.
- For columns 1-2, sample is the subset of HS6 products with trade in 2000 and 2005 for Chinese exports to US and aggregate EU-15. For columns 3-4, sample is the subset of HS6 products with trade in 2000 and 2005 for US imports from both Taiwan and China.
- Columns 1 and 3 differ only in the sample and the inclusion of the bilateral trade cost.

Table A4: Export growth from China: Robustness to HS-6 level and Pre-Accession Trends

<table>
<thead>
<tr>
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<tr>
<td>1</td>
<td>2</td>
<td>3</td>
</tr>
<tr>
<td>Uncertainty Pre-WTO (2000)</td>
<td>0.504** [0.223]</td>
<td>0.410* [0.225]</td>
</tr>
<tr>
<td>Uncertainty Pre-WTO (1996)</td>
<td>[-0]</td>
<td>[-]</td>
</tr>
<tr>
<td>Change in Transport Cost (Δln) ¹</td>
<td>-3.249*** [0.303]</td>
<td>-3.298*** [0.303]</td>
</tr>
<tr>
<td>Change in MFA quota status ¹</td>
<td>-0.378*** [0.112]</td>
<td>0.462*** [0.162]</td>
</tr>
<tr>
<td>Change in TTB status ¹</td>
<td>-0.118 [0.218]</td>
<td>-0.205 [0.306]</td>
</tr>
<tr>
<td>Observations</td>
<td>2,588</td>
<td>2,588</td>
</tr>
<tr>
<td>R-squared</td>
<td>0.047</td>
<td>0.052</td>
</tr>
</tbody>
</table>

Notes:
- Standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.1. Predicted sign of coefficient in brackets under variable.
- Robust regression employed to address potential outliers or influential individual observations due to double differencing. The estimation routine downweights outliers more than 7 times the median absolute deviation from the median residual. Uncertainty measure uses U.S. MFN and Column 2 Tariffs to construct profit loss measure at σ=3.
- (1) In columns 1 and 2 the change in tariff and transport cost variable represents double differences. In columns 3 and 4 they are single differences. Similarly for MFA and TTB variables.
### Table A5: Variety Growth from China (2000-2005)

<table>
<thead>
<tr>
<th></th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
</tr>
</thead>
<tbody>
<tr>
<td>Uncertainty Pre-WTO</td>
<td>0.253***</td>
<td>0.280***</td>
<td>0.240***</td>
<td>0.256***</td>
</tr>
<tr>
<td></td>
<td>[0.0731]</td>
<td>[0.0711]</td>
<td>[0.0890]</td>
<td>[0.0885]</td>
</tr>
<tr>
<td>Change in Tariff (Δln)</td>
<td>-2.680**</td>
<td>-0.729***</td>
<td>-2.263*</td>
<td>-0.733***</td>
</tr>
<tr>
<td></td>
<td>[-1.291]</td>
<td>[0.245]</td>
<td>[1.346]</td>
<td>[0.240]</td>
</tr>
<tr>
<td>Change in Transport cost (Δln)</td>
<td>-0.440***</td>
<td>-0.486***</td>
<td>-0.461***</td>
<td>-0.489***</td>
</tr>
<tr>
<td></td>
<td>[-0.165]</td>
<td>[0.163]</td>
<td>[0.162]</td>
<td>[0.160]</td>
</tr>
</tbody>
</table>

| Observations | 1,227        | 1,227        | 1,227        | 1,227        |
| R-squared    | 0.024        | .            | 0.061        | .            |
| Section FE   | no           | no           | yes          | yes          |
| Restriction p-value (F-test) | 0.12         | 1            | 0.246        | 1            |

Notes:
- Robust standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.10. Predicted sign of coefficient in brackets under variable.
- All specifications employ OLS and 2 and 4 impose theoretical constraint on tariffs and transport cost coefficients: \( b_{\tau} = b_d \sigma/(\sigma - 1) \). Constant included but not reported. The variety growth used as a dependent variable is measured by the ln change in the number of HS-10 products in a given each HS6. Sample: All regressions drop max-to-max transitions - observations at the maximum number of tradable HS-10 varieties at beginning and end of period - and zero-to-zero transitions that are not traded throughout the sample period.

### Figure A1: Panel Coefficients on Uncertainty Measure by Year

![Figure A1: Panel Coefficients on Uncertainty Measure by Year](image)

Notes: Results from an OLS unbalanced panel regression on log trade flows. Uncertainty measure in 2000 interacted by year. Coefficients are changes relative to the omitted group, the year 2000. Controls for applied tariffs, transport costs and dummy variables for section×year and HS-6 industry. Standard errors are clustered by HS-6. Two standard error bars plotted for each coefficient.
## Notation Reference

<table>
<thead>
<tr>
<th>Symbol</th>
<th>Description</th>
<th>Section</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\mu$</td>
<td>share of income spent on differentiated goods</td>
<td>2.1</td>
</tr>
<tr>
<td>$\Omega$</td>
<td>set of available differentiated goods</td>
<td>2.1</td>
</tr>
<tr>
<td>$E$</td>
<td>total expenditure on differentiated goods</td>
<td>2.1</td>
</tr>
<tr>
<td>$p_v$</td>
<td>consumer price of variety $v$</td>
<td>2.1</td>
</tr>
<tr>
<td>$\bar{P}$</td>
<td>price index for differentiated goods</td>
<td>2.1</td>
</tr>
<tr>
<td>$c_v$</td>
<td>unit labor cost for producer of variety $v$, the inverse of productivity $(1/c_v)$</td>
<td>2.1</td>
</tr>
<tr>
<td>$w_e$</td>
<td>wage in exporting country $e$</td>
<td>2.1</td>
</tr>
<tr>
<td>$d_V$</td>
<td>advalorem transport cost for industry $V$</td>
<td>2.1</td>
</tr>
<tr>
<td>$\pi(a_s,c_v)$</td>
<td>operating profits</td>
<td>2.1</td>
</tr>
<tr>
<td>$K, K_z$</td>
<td>sunk cost to start exporting or upgrading $(z)$</td>
<td>2.2, 2.6</td>
</tr>
<tr>
<td>$a_sV$</td>
<td>demand conditions for industry $V$ in state $s$: $a_sV \equiv A_{sV}^{-\sigma}d_{sV}^{-\sigma}$</td>
<td>2.2</td>
</tr>
<tr>
<td>$\beta$</td>
<td>probability that the firm/entrepreneur or consumer survive</td>
<td>2.2</td>
</tr>
<tr>
<td>$\Pi_e, \Pi$</td>
<td>expected value function of exporting $(e)$, and firm value function $\Pi$</td>
<td>2.2</td>
</tr>
<tr>
<td>$\lambda_s'$</td>
<td>transition probability from state $s$ to $s'$ of transition matrix $M$</td>
<td>2.3</td>
</tr>
<tr>
<td>$U_s'$</td>
<td>Uncertainty factor in state $s$ affecting entry and upgrade cutoffs</td>
<td>2.3</td>
</tr>
<tr>
<td>$\gamma$</td>
<td>policy uncertainty parameter, $\gamma \equiv 1 - \lambda_{11}$</td>
<td>2.3</td>
</tr>
<tr>
<td>$\omega$</td>
<td>Operating profit in col. 2 vs. MFN, partial equil.: $\omega \equiv (\tau_2/\tau_1)^{-\sigma}$</td>
<td>2.4</td>
</tr>
<tr>
<td>$u(\gamma)$</td>
<td>average spell a firm starting at $s = 1$ expects to spend in state $2$</td>
<td>2.4</td>
</tr>
<tr>
<td>$\lambda_2$</td>
<td>probability of state $s = 2$ conditional on exiting MFN state</td>
<td>2.4</td>
</tr>
<tr>
<td>$\ell$</td>
<td>labor endowment</td>
<td>2.5</td>
</tr>
<tr>
<td>$N$</td>
<td>mass of entrepreneurs</td>
<td>2.5</td>
</tr>
<tr>
<td>$\varepsilon_i$</td>
<td>elasticity of price index with respect to variable $i$</td>
<td>2.5</td>
</tr>
<tr>
<td>$\hat{\mu}$</td>
<td>parameters of indirect utility function: $\hat{\mu} = w_e \ell \mu (1 - \mu)^{(1-\mu)}$</td>
<td>2.5</td>
</tr>
<tr>
<td>$T$</td>
<td>time elapsed since transition from $s = 1$ to $2$</td>
<td>2.5</td>
</tr>
<tr>
<td>$\bar{\omega}$</td>
<td>Operating profit in col. 2 vs. MFN, general equil.: $\bar{\omega} \equiv (\tau_2/\tau_1)^{-\sigma}g$</td>
<td>2.5</td>
</tr>
<tr>
<td>$\zeta$</td>
<td>general equilibrium adjustment factor to profits lost in reversal</td>
<td>2.4</td>
</tr>
<tr>
<td>$\varepsilon'_V$</td>
<td>adjusted elasticity of price index with respect to $c_V$</td>
<td>2.4</td>
</tr>
<tr>
<td>$W_s$</td>
<td>consumer expected welfare at state $s$.</td>
<td>2.5</td>
</tr>
<tr>
<td>$n_m$</td>
<td>Expected number of periods under policy $m$ when starting at MFN.</td>
<td>2.5</td>
</tr>
<tr>
<td>$R_{sV}$</td>
<td>export level of industry $V$ in state $s$</td>
<td>2.5</td>
</tr>
<tr>
<td>$k$</td>
<td>shape parameter of the Pareto distribution for productivity $G_V(c)$</td>
<td>2.5</td>
</tr>
<tr>
<td>$\alpha_V$</td>
<td>industry specific distribution factor $\alpha_V \equiv \alpha_V \equiv N_{sV} \frac{\sigma}{c'_{sV}} \frac{k}{k - \sigma + 1}$</td>
<td>2.5</td>
</tr>
<tr>
<td>$\tilde{\alpha}_V$</td>
<td>industry modified factor in the export revenue $\tilde{\alpha}<em>V \equiv \alpha_V \left( \frac{1}{(1 - \beta)R</em>{sV}} \right)^{\frac{k - \sigma + 1}{\beta - 1}}$</td>
<td>2.5</td>
</tr>
<tr>
<td>$\zeta_V$</td>
<td>upgrading factor in exports for industry $V$, $\zeta_V \equiv 1 + \frac{R_{sV}}{K} \left( \phi_{sV} \right)^k &gt; 1$.</td>
<td>2.5</td>
</tr>
<tr>
<td>$f(U_V)$</td>
<td>general functional form for effect of uncertainty term on exports for industry $V$</td>
<td>3.1</td>
</tr>
<tr>
<td>$I$</td>
<td>tariff inclusive import penetration in total expenditure</td>
<td>C.1</td>
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</tbody>
</table>