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Policy Uncertainty, Trade and Welfare: Theory and Evidence for China and the U.S.

Kyle Handley
University of Michigan

Nuno Limão
University of Maryland and NBER

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Policy Uncertainty, Trade and Welfare: Theory and Evidence for China and the U.S.*

Kyle Handley
University of Michigan
handleyk@umich.edu

Nuno Limão
University of Maryland and NBER
Limao@umd.edu

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ABSTRACT: We examine the impact of policy uncertainty on trade, prices and real income through firm entry investments in general equilibrium. We estimate and quantify the impact of trade policy on China's export boom to the U.S. following its 2001 WTO accession. We find the accession reduced the U.S. threat of a trade war, which can account for over 1/3 of that export growth in 2000-2005. Reduced policy uncertainty lowered U.S. prices and increased its consumers' income by the equivalent of a 13 percentage point permanent tariff decrease. These findings provide evidence of large effects of policy uncertainty on economic activity and the importance of agreements for 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. For the U.S., that increase was even larger, rising from 3 to 19 percent. This has translated into a more than tenfold increase in the share of U.S. 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, which has led some authors to argue that the accession may have reduced trade costs faced by Chinese exporters.[1] But U.S. applied trade barriers toward China remained largely unchanged at that time.

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nation (MFN) status with accession, which ended the annual U.S. threat to impose high tariffs.[2] Had MFN status been revoked the U.S. would have reverted to Smoot-Hawley tariff levels and a trade war may have ensued.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."[3]

To examine this argument we build a model that allows us to interpret, measure and quantify 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, the welfare of U.S. consumers, and other outcomes. We focus on the role of TPU for investment and prices in part because of their importance in the context of the MFN debate. For example, the U.S. decision to delink MFN from China's human rights record was described as having "removed a major issue of uncertainty" and the renewal would have an impact on investment and re-exports that "will remove the threat of potential losses that would have arisen as a result of revocation."[4] U.S. business leaders argued that "...the imposition of conditions upon the renewal of MFN [was] virtually synonymous with outright revocation. Conditionality means uncertainty."[5] They lobbied Congress to make MFN permanent (Zeng, 2003). At the same time congressional research reports highlighted the higher consumer prices that would result if MFN was ever revoked (Pregelj, 2001). Our approach and results have important implications beyond this specific event; below we describe how they contribute to the growing literature on the impact of economic policy uncertainty and the role of trade agreements.

Our model captures the interaction between uncertainty and investment by modeling 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

[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.

[3] "China-U.S. trade volume increases 32 times in 23 years - Xinhua reports" BBC Summary of World Broadcasts, 2/18/2002.

[4] "HK business leaders laud U.S. decision" South China Morning Post, 5/28/94, Business section. The uncertainty recurred several times until the WTO accession.

[5] Tyco Toys CEO "China MFN Status," Hearing before the Committee on Finance, U.S. Senate, June 6, 1996, p. 97.

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. (2015) 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 and countries (capturing the differential 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 and differ in their sunk costs).

To guide the estimation and quantification we develop a dynamic heterogeneous firms model with TPU. We build on Handley and Limão (2015) and extend it in three ways. First, firms can invest not only to enter foreign markets but also to upgrade their export technology. This allows changes in uncertainty to affect the extensive margin (new exporters) and the intensive margin (continuing exporters with upgraded technology).[6] Second, the exporting country is allowed to be large enough to affect the importer's aggregate outcomes. Otherwise TPU has no significant impact on the importer. Third, entry into production is endogenous and subject to sunk entry costs such that TPU affects the formation and reallocation of firms. The model provides a number of insights. We highlight that TPU has both a direct and indirect effect on firm outcomes. The direct effect of TPU is to lower entry through an option value of waiting for exporters (fear of higher protection) and domestic firms (fear of low protection). The effect of these entry reductions is to increase the price index of the importer, which is central to the welfare gains from reforms that lower TPU. This price index increase has an indirect positive effect on exporter and domestic entry that can dominate for exporters (if initial protection is already very high) or for domestic firms (if initial protection is already low).

As preliminary evidence and motivation for why we require a theoretical framework, consider Figure 2. In panel (a) we plot Chinese average export growth to the U.S. between 2000-2005 by sector against the (log) difference of the column 2 and MFN tariffs in 2000. On average, those sectors facing a relatively higher initial tariff threat in the case of MFN revocation experienced faster export growth and larger declines in prices, as shown in panel (b). The exercise is suggestive, but also raises questions regarding the identification of partial effects and the quantification of the general equilibrium effects, both of which the model helps to address. First, what is a theory-consistent measure of uncertainty? The model shows it is the proportion of profits that Chinese exporters would lose if China ever lost its MFN status. We map this to observable tariff measures and then find evidence that our measure is relevant to exporters. Second, what are the necessary controls and assumptions required to identify the TPU effect and what structural parameters can

[6] Evidence for both margins in China's export boom is documented by Amiti and Freund (2010), 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).

we estimate? The model generates a tractable TPU-augmented gravity equation that allows us to consistently aggregate individual firm decisions to the industry level and identify the change in the probability of MFN being revoked. Moreover, the model generates a relationship between ideal import price indices and TPU that we also estimate. Third, the model predicts these effects should only apply to trading partners where TPU changed and in industries with sunk costs of exporting.

We use variation in policies, export values and prices across thousands of products to estimate the effects of TPU. We find non-parametric and parametric evidence that Chinese export growth in 2000-2005 was higher in industries with higher initial TPU. The effect is robust to controlling for applied tariff and non-tariff barriers, transport costs and sector specific growth trends. The effect is only present in industries with export sunk costs, which we identify by exploring persistence in export behavior. Moreover, the effect is also robust to allowing for a broader set of shocks than those present in the theoretical model; namely unobserved shocks to import demand (TPU has no direct effect on other U.S. imports) and export supply (U.S. TPU toward China has no direct effect on Chinese exports to non-U.S. destinations), which rules out a large set of potential confounding factors.

We also construct industry level ideal import price indices following Feenstra (1994) and find larger reductions in industries with initially higher TPU. This is the effect the model predicts due to new imported varieties (for which we find direct evidence) and technology upgrading. The price effect is also robust to controlling for alternative variables and unobserved import demand shocks and it is only present in high sunk cost industries. The partial effect of reducing TPU was to lower the average U.S. industry price indices for Chinese imports by at least 15 log points and the corresponding aggregate index by slightly more.

The significant partial effects of TPU on import prices leads us to quantify its aggregate effects. In section 4 we characterize the general equilibrium effects of TPU by solving for the model in changes. We derive the impacts on firm entry, sales and prices (foreign and domestic) and how they depend on key features of the policy regime: current and future tariffs and the probability of transitioning between them. Combining this framework with a non-linear estimate of the TPU-augmented gravity equation we identify the reduction in the probability of MFN revocation. To isolate and quantify the aggregate effects of reducing TPU we then evaluate the impacts of the estimated shock to this structural parameter. The counterfactual implies an aggregate Chinese export increase of 32 log points, which is about one third of the observed growth in this period. The predicted changes in the U.S. import price index, domestic manufacturing firm sales, employment and entry are also consistent with the observed changes during this period. The counterfactual import penetration if TPU had remained in place between 2002-2010 would have been substantially lower, as shown by the dashed line in Figure 1.

We also contribute to the long standing question of the aggregate gains from trade. Recent work by Arkolakis et al. (2012) shows that import penetration and trade cost elasticities are sufficient statistics to compute those gains in a class of models. That is also the case for the deterministic version of our model, and so the gains from trade, or autarky cost, provides a useful benchmark. However, under TPU those are no longer sufficient statistics and we require the change in the ideal price index. We estimate that TPU increased that U.S. price index (for tradeables) by half as much as fully eliminating trade with China. So the U.S. consumer welfare cost of TPU was about half that of going to autarky, or the equivalent of permanent tariff increase of 13 percentage points on Chinese goods.

Understanding the impact of TPU has broader implications beyond this episode. It informs us about the potential impacts of other sources of policy uncertainty, such as U.S. threats to impose tariffs against "currency manipulators" 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, until recently, was largely missing from the analysis of trade agreements. 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, 2013). Moreover, the welfare cost of trade wars in the absence of such agreements are potentially large—about 3.4% of income according to some quantitative exercises (Ossa, 2014). 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 (2015) 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 direct evidence for welfare gains from reducing TPU through trade agreements. Finally, we illustrate how the model applies beyond the Chinese accession through various counterfactual exercises where the U.S. unilaterally abandons all its trade agreement commitments and increases TPU and/or applied tariffs.

Our research also complements the recent empirical work on the impact of Chinese exports on developed countries. Bloom et al. (2011) assess the impact of Chinese exports on wages and employment in the European Union while Acemoglu et al. (Forthcoming) and Caliendo et al. (2015) focus on the U.S. Pierce and Schott (Forthcoming) study the effects of Chinese exports on U.S. manufacturing employment and, as an intermediate step, they estimate the reduced form effect of column 2 tariffs on exports.[7] Our papers differ in important ways. First, our focus is on the trade, price and consumer welfare effects. Second, we provide evidence for the central mechanism: sunk costs of exporting. Third, we develop a theoretical framework that contributes to the literature on agreements and gains from trade while allowing for the structural identification of parameters. Among other things, we explore the counterfactual exercises to isolate and quantify the aggregate effects of TPU on several outcomes and decompose them, e.g. we find that a large fraction of the trade and price changes is explained by a mean preserving compression of the tariff and the rest is due to locking in tariffs below the mean.[8]

We present the basic framework and derive the TPU-augmented gravity equation in section 2, followed by the empirical analysis in section 3. The general equilibrium solution in section 4 is used for the structural estimation and quantification in section 5. The appendices contain details on the theoretical derivations and empirical implementation.

[7]Therefore, independently from us, they too follow the proposal in Handley and Limão (2012) to estimate the importance of "the U.S. threat of non-renewal of China's MFN status and whether its elimination in 2001 (upon China's WTO entry) can explain the subsequent export boom to the U.S." (p. 44).

[8]In a working paper version we also quantify the uncertainty impact of proposed legislation that threatens to impose tariffs of almost 30% on "currency manipulators". We 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.

2 Framework and Partial Equilibrium Effects

We first describe the basic framework and firm entry decision problems, which apply throughout the paper. We then derive the effect of TPU on these decisions from the perspective of a small exporting country—one that takes foreign aggregate variables as given. We initially focus on a single industry and, in section 2.5, we model multiple industries and technology upgrading, which we use to derive the TPU-augmented gravity equation. This partial equilibrium structure is sufficient to derive and empirically identify any effect of TPU on exports. But in order to quantify its effects on exporter and importer outcomes, we allow for a large exporter and endogenous domestic entry in section 4.

2.1 Demand, Supply and Pricing

Consumers spend a fixed share of 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 $P = \left[\int_{v\in\Omega} (p_v)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}$ is the CES price index for the set of available varieties, Ω.

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 underlying productivity and the distribution of other firms in each market.

The consumer price, p_v, includes an ad valorem tariff, $\tau \geq 1$, so exporters receive p_v/τ per unit (domestic producers face no taxes in their market). The tariff is common to all firms in the differentiated industry. After observing τ each firm chooses p_v to maximize operating profits taking aggregate conditions as given, and correctly anticipating their equilibrium value. We allow for an ad valorem 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

$$\pi(a, c_v) = ac_v^{1-\sigma} \tag{1}$$

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

2.2 Policy Uncertainty and Entry

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 current realization of the policy, and; (iii) 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 period 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.

For any state s the expected value from exporting for any firm v after entry is

$$\Pi_e\left(a_s, c\right) = \pi\left(a_s, c\right) + \mathbb{E}_s \sum_{t=1}^{\infty} \beta^t \pi\left(a'_s, c\right) \tag{2}$$

where we omit the variety subscript for simplicity; $\mathbb{E}_s$ 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 enters if its cost is below a threshold value, c_s^D. This benchmark threshold is obtained by equating the present discounted value of profits to the sunk cost.[9]

$$\frac{\pi\left(a_s, c_s^D\right)}{1-\beta} = K \Leftrightarrow c_s^D = \left[\frac{a_s}{(1-\beta)K}\right]^{\frac{1}{\sigma-1}} \tag{3}$$

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

$$\Pi(a_s, c) = \max\left\{\Pi_e(a_s, c) - K, \beta\mathbb{E}_s\Pi(a'_s, c)\right\}. \tag{4}$$

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

$$\Pi(a_s, c_s^U) = \Pi_e(a_s, c_s^U) - K, \tag{5}$$

and any firms with lower costs will enter in state s.

Production entry

It will be clear that our estimation strategy for the partial effect of TPU on exports is valid under alternative assumptions regarding entry into production. However, the general equilibrium effects of TPU will depend on production entry decisions. We model the latter similarly to export entry: to start production, a firm requires a sunk cost, K_h, in order to activate a known technology. The firms make this decision after observing the current realization of the policy. Thus firms with a cost below a certain threshold enter and continue to produce the following period with a fixed probability, their survival rate (there are no production fixed costs). The domestic operating profit of a home firm is $\pi_h = a_h c_v^{1-\sigma}$ where $a_h = (\sigma)^{-\sigma}\left((\sigma-1)P\right)^{\sigma-1}E$ and we assume that $K_h \geq 0$ is sufficiently small that the marginal home entrant does not export. Therefore

[9] This is an implicit solution for the cutoff if exporters are large since then a_s depends on c_s^D.

the domestic entry thresholds for home market firms can be obtained using the expressions we derived above when evaluated at K_h, $a_{s,h}$ and β_h; specifically by using (3) if there is no uncertainty to obtain $c_{s,h}^D$ and (5) to determine $c_{s,h}^U$.

2.3 Policy Regime

To characterize the effects of TPU we propose an exogenous policy process that captures three key states of trade policy, denoted $m = 0, 1, 2$. Standard models of trade policy consider permanently high or low protection states, where $\tau_2 > \tau_0$. These extremes can capture outcomes under no cooperation (e.g. U.S. tariffs on Cuba or North Korea) or under a credible agreement (e.g. U.S. tariffs on Canada or certain WTO members). To analyze the effect of TPU we add an intermediate protection state characterized by a temporary tariff $\tau_1 \in [\tau_0, \tau_2]$ that changes with probability γ. Formally, the **trade policy regime** is characterized by a Markov process with time invariant distribution, denoted by $\Lambda(\tau_m, \gamma)$.

By allowing for three states we can capture a rich set of situations.[10] To address the central questions we can focus on a simple transition matrix where policy is uncertain only in the intermediate state, so $\gamma > 0$, and the extreme states are absorbing.[11] The exact interpretation of each state depends on the setting. In our empirical application the intermediate state captures China's pre-WTO period when its temporary MFN status in the U.S. could change with probability γ and give way to either high protection (column 2 tariffs) with probability λ_2, or low protection with probability $1 - \lambda_2$. So we can interpret WTO accession as a switch to the low state in this application. Alternatively, we can interpret the WTO accession as an exogenous change in policy regime, if it lead to an unanticipated change in γ (or τ_m). Thus we derive the effects of transitions across policy states within a regime and transitions across regimes.

The three state process has two other benefits when considering countries starting in the intermediate protection state. First, it captures the two key effects of agreements: reducing applied protection and/or TPU. Second, it allows for the possibility that policy worsens for either foreign firms (higher protection state) or domestic ones (lower protection state); this generates an option value of waiting for both types of firms. These benefits will become clearer in section 4 when we decompose and quantify the applied and TPU effects and account for the general equilibrium responses of domestic firms.

2.4 Partial Equilibrium

To estimate the impact of trade policy on entry and exports we derive the effects on cutoffs from switching regimes or switching states and decompose the latter into a change in applied policy and policy uncertainty.

Tariffs are the only underlying source of uncertainty and we initially focus on a small exporting country such that changes in its exports have a negligible effect on the importer's aggregate variables, E and P. In the estimation we control for any unexpected shocks to aggregate variables.

The small country assumption implies that tariff changes only have a direct impact on market conditions for the exporters, and there is one distinct value of a_s for each τ_m for any given value of $EP^{\sigma-1}$.[12] In

[10] These include any setting where there is some probability of (i) cooperation with negligible probability of increases in protection (e.g. under credible agreements) (ii) partial cooperation (e.g. when protection may increase but a credible agreement is also possible) and (iii) higher protection levels (including a trade war or even autarky).

[11] In the appendix we argue this is a special case of a less restrictive requirement, that $\Lambda(\tau_{m+1}, \gamma)$ first order stochastically dominates $\Lambda(\tau_m, \gamma)$ for $m = 0, 1$, and show how key results generalize if the high state is not absorbing.

[12] We do not need to impose additional general equilibrium structure to solve for these export cutoffs. Uncertainty changes

appendix A.1 we use this to show that the solution to the Bellman equation (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_s^U, for each τ_m. The cutoff for the intermediate state, c_1^U, is proportional to its deterministic counterpart in (3) by an **uncertainty factor**, $U(\omega,\gamma)$.

$$c_1^U/c_1^D = U(\omega,\gamma) \tag{6}$$

$$U(\omega,\gamma) \equiv \left[\frac{1+u(\gamma)\,\omega}{1+u(\gamma)}\right]^{\frac{1}{\sigma-1}} \tag{7}$$

If U is less than one, then entry is reduced 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 the intermediate state (given no other conditions changed). The term $u(\gamma) \equiv \gamma\lambda_2\frac{\beta}{1-\beta}$ is the average spell that a firm starting at the intermediate state expects to spend under τ_2. This spell is increasing in the probabilities of: exiting the intermediate policy state, γ, and then facing a higher tariff, λ_2, and surviving, β. Note that if $\gamma = 0$ then policy is fixed in all states, thus we say that there is **policy uncertainty** if $\gamma > 0$. Moreover, any increase in γ implies a higher probability of a policy change but does not change the odds of the worst or best case scenario. We interpret this as an increase in policy uncertainty.[13]

From these expressions we can see that entry in the intermediate state is lower under uncertainty if and only if tariff increases are possible, i.e. $c_1^U < c_1^D$ iff $\tau_2 > \tau_1$ and $u(\gamma) > 0$. Note that while TPU can lead to lower or higher tariffs, it is only the possibility of high tariffs that affects export entry; if there is uncertainty, $\gamma > 0$, but tariff increases are not possible, $\lambda_2 = 0$, then uncertainty has no impact on entry.

The entry result in (6) reflects a specific switch in policy regime: an unanticipated introduction of TPU at a given tariff. We note two simple extensions that are relevant for the empirical analysis. First, the effect of TPU on entry is monotonic ($dc_1^U/d\gamma < 0$ for all γ) so we can also test for marginal changes in TPU, e.g. whether before WTO accession Chinese exporters faced higher TPU in years when an MFN revocation seemed more likely. Second, we also want to understand the effect of agreements that are anticipated with some probability, i.e. switches to state 0, and compare them to unanticipated changes in TPU. In the appendix we show the cutoff in the intermediate state relative to any deterministic baseline state with tariffs τ_b is

$$c_1^U/c_b^D = U(\omega,\gamma) \times (\tau_1/\tau_b)^{-\frac{\sigma}{\sigma-1}}\,. \tag{8}$$

If $\tau_b = \tau_0$ this expression captures the reduction in applied policy and uncertainty from entering state 0 since, when there is no uncertainty in that state, the cutoff is equal to the deterministic value, $c_0^U = c_0^D$. Switching from the intermediate to the low state increases entry by reducing applied tariffs by τ_1/τ_0 and/or TPU. Thus, even if an agreement is anticipated with some probability, entering it can be used to identify an unanticipated elimination of TPU after we control for applied tariff changes.

The impact of eliminating TPU, as we have defined it, can be decomposed into a pure risk and expected mean effect. To understand each of these consider the regime switch described above when we start in the intermediate state and uncertainty is eliminated. If τ_1 was at the long-run mean of the original tariff process then this uncertainty reduction is exactly a mean preserving compression of tariffs, or a *pure risk* reduction.[14]

the profits from exporting but these are separable from domestic profits since the wage is unity in any diversified equilibrium.

[13]This implies $\frac{\lambda_{12}}{\lambda_{10}} = \frac{\gamma\lambda_2}{\gamma(1-\lambda_2)}$ where λ_2 is the probability of higher tariffs conditional on exiting the intermediate state.

[14]In this three state process, when state 1 has a policy τ_1 equal to the long-run mean then a decrease in γ induces a mean

However, if τ_1 was below its long-run mean, as will be the case in our application, then the reduction in γ has the additional effect of locking in tariffs below their expected value under uncertainty. The structure of the model will help us quantify each of these effects.

In sum, the explicit solution for the entry cutoff in eq. (6) allows us to derive its elasticity with respect to γ, and the appropriate measure to capture the potential losses under the worst case scenario, ω. Next we show how to explore variation in this measure over industries to identify the effect of TPU.

2.5 TPU augmented gravity

We derive a TPU augmented gravity equation to estimate how changes in policy uncertainty translate into export growth. This requires extending the baseline model in two dimensions. First, we model the effect of uncertainty on the intensive margin of a firm's exports. Second, we allow for industry variation in policies, which is necessary for our identification strategy.

Technology Upgrade

We will focus on estimating the effect of changes in TPU on export growth. This growth can reflect extensive and intensive margin effects and we now show how the TPU augmented gravity can capture both. We believe this extends the applicability of the framework to situations where both margins are potentially important. For example, most Chinese export growth to the U.S. in 2000-2005 took place in HS-10 goods that were already being exported in 2000. Some of the growth in existing products is due to new exporters but it is also plausible that existing ones grew by investing in export activities due to reduced TPU in the U.S. We model one potential channel, irreversible investments by incumbent exporters to upgrade their technologies. This is consistent with the large increases in TFP growth of Chinese firms since WTO accession.[15]

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.[16] The operating profits are then $\pi_v = a_s \left(zc_v\right)^{1-\sigma}$. In the Online Appendix C.2 we show that the upgrading decision is similar to the entry decision in that it also takes the form of a cutoff cost. The upgrade cutoff is $c_{sz}^U = \phi c_s^U$. It is proportional to the entry cutoff by a constant *upgrading parameter*, $\phi \equiv \left[\left(z^{1-\sigma} - 1\right) \frac{K}{K_z}\right]^{\frac{1}{\sigma-1}}$. 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 export entry cutoff solutions will be similar to those we derived, but only the more productive exporters will upgrade. Since ϕ is independent of tariffs, the elasticity of the upgrade and entry cutoff are exactly the same with respect to tariffs and uncertainty.

Multi-industry Aggregation

preserving compression of the initial conditional policy distribution, $\Lambda\left(\tau_1, \gamma\right)$. This is one motive to use a 3-state process.

[15] 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. In future work we plan to directly estimate if there is a causal effect of TPU on Chinese firm TFP.

[16] An interpretation of the ad valorem 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 it.

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. It thus captures our empirical setting such that if any industry V moved to the agreement state (or the high protection state) then all industries would face the same policy state.[17]

The export revenue in state s for firm $v \in V$ is $p_{sv}q_{sv}/\tau_{sV} = a_{sV}\sigma(z_V c_v)^{1-\sigma}$ once we plug in the optimal price and quantity from section 2.1, where $z_V < 1$ for the upgraders and unity otherwise. The economic conditions variable, a_{sV}, still reflects aggregate income and price index effects but it now reflects industry specific tariffs (and export costs). The export entry cutoff in eq. (6) is industry specific but otherwise the previous cutoff results are unchanged.

The mass of exporting firms in any stationary equilibrium, characterized by a constant mass of active firms, is equal to $N_V \times G_V(c_s^U)$, the total number of potential firms in industry V in the export country times the fraction of these with costs below the cutoff. Export revenue in industry V is obtained by aggregating over firms that upgrade and those that do not as follows[18]:

$$R_{sV} = a_{sV} d_V^{1-\sigma}\sigma N_V \left[\int_0^{\phi_V c_{sV}^U} (z_V c)^{1-\sigma}\, \mathrm{d}G_V(c) + \int_{\phi_V c_{sV}^U}^{c_{sV}^U} c^{1-\sigma}\mathrm{d}G_V(c)\right] \tag{9}$$

We assume that productivity in each industry is drawn from a Pareto distribution bounded below at $1/c_V$, so $G_V(c) = (\frac{c}{c_V})^k$ 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 (9), use the definition of a_{sV}, and c_{sV}^U, and take logs to obtain an uncertainty augmented gravity equation,

$$\ln R_{sV} = (k - \sigma + 1)\ln U_{sV} - \frac{\sigma}{\sigma - 1} k \ln \tau_{sV} - k \ln d_V + k \ln P + \frac{k}{\sigma - 1}\ln E + \ln \zeta_V + \ln \tilde{\alpha}_V \ . \tag{10}$$

Without either policy uncertainty, $U_{sV} = 1$, or upgrading, $\zeta_V = 1$, (10) reduces to a standard industry level gravity equation conditional on aggregate expenditure on differentiated goods, E, and the price, P (cf. Chaney, 2008). The terms $\tilde{\alpha}_V$ and ζ_V are combinations of industry parameters that are time invariant.[19] Finally, all else equal, upgrading increases export levels, as reflected in ζ_V, but not the *elasticity* of industry exports with respect to U_{sV}. 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.

3 Estimation and Identification

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. The objective of this section is to identify a causal effect of TPU on exports and test if the data is consistent with some of the model's assumptions. We use the TPU augmented

[17] Industries can have different tariff levels, as long as their ranking across states is the same.

[18] We use the relationship $c_{sz}^U = \phi c_s^U$ and allow industry variation in the upgrade technology and sunk costs. Away from stationary equilibria there are additional exporters who entered in previous periods under better conditions.

[19] $\tilde{\alpha}_V \equiv \frac{N_V \sigma}{c_V^k}\frac{k}{k-\sigma+1}\left(\frac{1}{(1-\beta)K_V}\right)^{\frac{k-\sigma+1}{\sigma-1}}\sigma^{\frac{-\sigma k}{\sigma-1}}(\sigma-1)^k$ and $\zeta_V \equiv 1 + \frac{K_{zV}}{K_V}(\phi_V)^k$

gravity equation in (10). As we show in section 4, this same equation holds even when the exporter is large enough to affect the aggregate variables. If we control for those variables, then we can identify whether TPU affected exports regardless of exporter size.

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 ζ_V.[20] 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 uses the time-difference of (10),

$$\Delta \ln R_V = f\left(\frac{\tau_{2V}}{\tau_{1V}}, \gamma\right) + b_\tau \Delta \ln \tau_V + b_d \Delta \ln D_V + b + e_V \tag{11}$$

where $\Delta \ln$ represents the difference between post- and pre-WTO periods. The coefficient $b_\tau = -\frac{k\sigma}{\sigma-1} < 0$ is the effect of applied tariffs (conditional on the uncertainty factor). We model ad valorem export costs, d_V, as a function of observable shocks given by the ad valorem 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 costs 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 null hypothesis under the model is that prior to WTO accession there is a positive probability of column 2 tariffs, i.e. $\gamma\lambda_2 > 0$, and thus f is increasing in τ_{2V}/τ_{1V} if the accession reduced γ. If the accession eliminated uncertainty then $f = -(k - \sigma + 1) \ln U\left(\frac{\tau_{2V}}{\tau_{1V}}, \gamma\right)$.[21]

Standard trade models with a gravity structure yield a restricted version of (11) 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 τ_{2V}/τ_{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 and unobserved demand and supply shocks). Fourth, after introducing the general equilibrium structure we provide a non-linear structural estimate of the corresponding term for f that we then use to quantify the impact of TPU.

Importantly, we also examine if the uncertainty effect is present only in industries where sunk costs are important, as the model predicts.

[20] We address the possibility that these terms are time varying and correlated with TPU in the robustness section.

[21] This is the case whether entry is interpreted as a switch in state or as an exogenous change in regime, as discussed in section 2.4. Under the regime switch interpretation the model allows for $\gamma > 0$ even after entry. In that case $f = (k - \sigma + 1) \Delta \ln U\left(\frac{\tau_{2V}}{\tau_{1V}}, \gamma\right)$ where changes in U reflect either changes in tariffs or γ. In the period we consider τ_{2V}/τ_{1V} is nearly constant within industries, so f captures changes in γ. So in this case f is also increasing in τ_2/τ_1.

3.2 Data and Policy Background

We combine trade and policy data from several sources. U.S. import trade data at the HS-10 level for several years is obtained from the NBER Harmonized System Imports by Commodity and Country. These data are concorded over time and aggregated to the HS-6 level for the export growth analysis at the HS-10 level to compute ideal price indices and examine the persistence of exporting. We obtained U.S. tariff schedules via the World Bank's WITS. The source of other policy measures we use are described in Appendix B. The cost of insurance and freight is reflected in the import data. We convert all tariff and transport cost data to their iceberg form (e.g. from 10% to $\ln \tau = \ln 1.1$).

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.[22]

Finally, we highlight some useful policy background for understanding the timing choice for our baseline estimates. Uncertainty remained about both China's accession to the WTO and its permanent normal trade relations (PNTR) as late as 2000 due to tense foreign and economic relations.[23] As a result, protracted negotiations over China's accession meant Congress voted again in the summer of 2001 over whether to revoke MFN. China joined the WTO on December 11, 2001 and the U.S. effectively enacted PNTR on January 1, 2002. 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 Partial Effect Estimates of TPU on Exports

Table 1 provides summary statistics for our baseline sample. Export growth from 2000 to 2005 averaged 128 log points (lp) across HS-6 industries, with substantial variation across them: the standard deviation is 168 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.

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 σ the industry ranking of potential profit loss is determined by τ_{2V}/τ_{1V} so we use this ratio to partition the sample into the columns in Table 1 labeled low (bottom tercile of τ_{2V}/τ_{1V}) and high TPU industries. Export growth in high uncertainty industries is 19 lp higher, a mean difference that is statistically significant. The export growth distribution for high TPU industries first order stochastically dominates the one for low TPU industries, as shown by the respective kernel densities in figure 3(a) and confirmed via a Kolmogorov-Smirnov test. Figure 4(a) provides further non-parametric evidence of this relationship by estimating a local polynomial regression of export growth on $\ln(\tau_{2V}/\tau_{1V})$. We confirm the higher growth in high initial uncertainty industries, as obtained in the mean test, and a non-negative relationship over the

[22] Our baseline sample is smaller because it focuses on the 94% of HS-6 lines where tariffs are levied on an ad valorem basis, some have only specific tariffs. We show the results are robust to this and the zero trade flow industries in section 3.4.

[23] 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. 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.

full domain.

Semi-parametric evidence and functional form

Using a semi-parametric approach we can control for other determinants of export growth and test for specific functional forms of the uncertainty term. Several trade models yield a gravity equation that is a special case of (11) with the implicit restriction that $f = 0$. We use the residuals from that restricted estimation to determine how τ_{2V}/τ_{1V} affects f without imposing functional forms. Using a double residual semi-parametric regression (Robinson, 1988) we find that τ_{2V}/τ_{1V} 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 heterogeneous growth trends in sectors.

In Figure 4(c), we plot the resulting semi-parametric fit (that did not impose any σ) and see it is increasing in $1-\left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-3}$ — the potential profit loss measure when $\sigma = 3$. The predicted parametric line, obtained from OLS estimation of (11) using this parametric loss measure lies everywhere within the semi-parametric 95% confidence interval. We fail to reject the equality of this particular parametric fit and the semi-parametric one and will thus use OLS specifications as a baseline from which to perform robustness tests. We choose $\sigma = 3$ and test if results are robust to alternative values. We also test if the semi-parametric fit is equal to alternative parametric fits that are linear or log linear in τ_{2V}/τ_{1V} 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 τ_{2V} is smaller at high tariffs where trade would be negligible.

Parametric OLS estimates

The semi-parametric evidence supports approximating the uncertainty term using $b_\gamma \times \left(1-\left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-3}\right)$ in (11). When we approximate U_V linearly around $\gamma = 0$ and use (7) we have the following structural interpretation of $b_\gamma^{OLS} = \frac{k-\sigma+1}{\sigma-1} u(\gamma)\, g \geq 0$.[24] We first present parametric estimates of b_γ and check 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_γ 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 cannot reject. We therefore re-estimate the model in column 2 with this restriction, which increases the precision of the estimates.[25]

Sector level growth trends and unobserved heterogeneity: The model contains several unobservables that can vary across industries. Most of unobservables are time invariant and log separable and are differenced out in the baseline estimates, e.g. sunk costs, upgrade technology and the mass of non-exporting Chinese firms. Any growth innovations common to all industries are absorbed in the baseline constant, b. We now allow for that growth to differ across 21 sectors by including a set of dummies in the differenced equation

[24] If the exporter is small then $g = 1$ but if it is large then g will be slightly different from unity as shown in section 4.

[25] 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.

(11). The results in columns 3 and 4 of Table 2 are similar to those in the baseline—importantly, b_γ 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, upgrading parameters, FDI and Chinese barriers on intermediates. In subsequent results we will control for even finer unobserved shocks at the industry level.[26]

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.[27] In column 2 of Table 3 we control for the change in the binary indicator for both MFA quotas and NTBs and find they have the expected negative sign.[28] 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. 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 earlier 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).[29]

3.4 Robustness

Elasticity of substitution, outliers and sample selection

Table A2 summarizes the robustness of the baseline linear estimates of b_γ (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_γ 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; 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

[26]There are 21 sectors (or sections) including Machinery, Textiles, Footwear and 18 other groups that the UN defines as coherent groups of HS-2 categories. In figure 2(a) we plot the average export growth on the mean of $\ln(\tau_{2V}/\tau_{1V})$ by sector and find a positive correlation. If most of the variation in TPU or export growth was across sectors then we would worry their relationship would be proxying for an omitted sector effect. However, the summary statistics in Table A1 show there is considerable variation in both variables within sectors.

[27]Additional details on the NTB and MFA indicators and sources appear in the data Appendix B.

[28]The yearly panel evidence in section 3.5 shows 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. Those panel estimates are robust to dropping products that ever had an MFA quota regardless of the year it was removed.

[29]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).

have elasticities very different from the overall median we do the following. Let $\widehat{\sigma}_V$ denote the median HS-10 elasticity estimate (from Broda and Weinstein) in each HS-6 industry. In columns 5 and 6 we use $\widehat{\sigma}_V$ directly to recompute the uncertainty measure and obtain similar results.[30]

Potential outliers. In columns 1 and 2 of Panel B we employ a robust regression procedure that down-weights 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 the log change, which reduces the sample size. In columns 3 and 4 of Panel B we address this by using the mid-point growth rate, which can accommodate zeroes, as our dependent variable. The baseline also excludes industries that only have specific tariffs. Columns 5 and 6 add these additional industries by calculating ad valorem equivalents (given by specific $\text{tariff}_V/\text{unit value}_V$) and incorporating them into both the change in applied tariffs and the uncertainty measure.

Processing trade. Chinese exports in certain industries primarily reflect 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 processing industries, then they could reflect changes in Chinese policies towards processing trade. In columns 7 and 8 of Panel B we drop all the HS-6 industries section XVI of the Harmonized System, which has the largest share of processing trade. Below we show the results are robust to other unobserved supply shocks.

Unobserved Supply and Demand Shocks

We addressed omitted variable bias thus far by controlling for specific variables at the HS-6 level and unobserved contemporaneous sector shocks. There are other industry shocks one could consider, e.g. differential growth depending on labor intensity.[31] In Tables 4 and 5 we go one step further and provide evidence that the baseline results are robust to controlling for unobserved HS-6 industry demand and supply shocks.

Unobserved 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 the shock was specific to China 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 Table 4 by regressing Chinese export growth to the European Union and Japan on the U.S. TPU measure. Whereas for the U.S. we found a positive and significant effect of this measure (column 1), we do not find significant effects for Chinese exports to the E.U. or Japan (columns 2 and 3 respectively).[32] In column 4 we pool all three samples, include a full set of HS-6 effects, and cluster standard errors at the HS-6 level. Thus we control for narrowly defined Chinese supply shocks, including any Chinese policy changes induced by WTO accession and technology changes that are not export market specific. The coefficient on the TPU measure remains positive and significant and now identifies the average differential growth effect of U.S. TPU on Chinese exports to the U.S. relative to the E.U. or Japan in the same industry.

[30] As we can see from the estimating equation when there is variation in σ_V the parameters are not constant so we obtain an average effect. We can take this into account by estimating different coefficients for the tariff and uncertainty variables, one for each tercile of $\widehat{\sigma}_V$, doing so we can't reject the equality of those coefficients. We obtain similar results if we instead assume $\sigma = 3$ but drop any industry with $\widehat{\sigma}_V \notin [1.5, 4.5]$.

[31] We show the results are robust to controlling for a measure of labor intensity in Table A5.

[32] Column 1 shows this specification for the U.S. applied to the common subsample of industries exporting to both destinations. We do not include the transport cost since we do not have that data for the E.U. and Japan.

Unobserved demand shocks. If U.S. production decreased (and/or its consumption increased) in industries where China faced higher initial uncertainty, then the baseline estimates could be biased upward. But such shocks would also increase U.S. imports from other countries. We do not find support for this. In Table 5 we pool U.S. imports from China and other countries that in 2005 faced the same policy regime as China, i.e. WTO members without a U.S. PTA.[33] In column 1 we estimate a positive and significant effect of U.S. TPU on Chinese imports and no significant effect on non-Chinese imports. We cluster by HS-6 industry because we have no variation in applied tariffs or the TPU measure by country. In column (2) we include an additional set of HS-6 digit industry effects to control for any U.S. demand or production shocks. These industry effects also control for any unobserved change in industry trade barriers and observable MFN tariffs, which are no longer identified. The differential effect on Chinese imports remains positive and significant. In columns 3 and 4 we focus solely on Taiwan and China, which allows us to keep a number of other important factors constant. First, Taiwan also acceded to the WTO in January 2002, right after China. 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. The results in column 3, with sector-country effects, and column 4, which adds HS-6 dummies, support this prediction.

3.5 Additional Evidence: Sunk Costs and Timing

Sunk Cost Channel

Our model predicts that TPU reduces trade when sunk costs are present and there is an option value of waiting for some firms. In our baseline regressions we control for any time invariant HS-6 industry variation in sunk costs. In the falsification tests above we further control for any changes in those costs. In both we obtained an average TPU elasticity across all industries. We now test if that elasticity is higher for high sunk cost industries. Doing so requires a measure of sunk costs of exporting and, since none is readily available, we estimate it by exploiting variation in the persistence of exporting over time. A standard approach (cf. Roberts and Tybout, 1997) is to use firm-level data to estimate a probability model where, after conditioning on firm characteristics to capture their current incentive to participate, any correlation with lagged participation provides evidence of sunk costs. We apply a similar approach but use HS-10 product data and estimate persistence parameters, b_V^{sunk}, industry-by-industry. To minimize the impact of the episode we study on the persistence estimates we use data for the pre-accession period, 1996-2000, and focus on the effect of lagged export participation on current export participation for each HS-10 product-country variety for all U.S. non-preferential trade partners, excluding China. In Appendix B.3 we argue that the estimates appear reasonable and provide additional details on the estimation and identification. We classify industries as having relatively higher sunk costs if their t-statistic for b_V^{sunk} is higher, which indicates we can be more confident of rejecting $b_V^{sunk} = 0$. More specifically, the top two terciles of the regression sample with higher t-statistics are classified as higher sunk cost relative to those in the bottom tercile.

In Panel A of Table 6 we estimate the effect of uncertainty across low and high sunk costs as we just defined. We find no effect of TPU on Chinese exports in low sunk cost industries, in column 1, in contrast we find a positive and significant effect for the high sunk cost subsample, in column 2. This provides strong

[33] We only use the non-preferential trade partners that face the same U.S. applied MFN tariffs as China.

evidence for the channel highlighted by the model's option value of waiting framework.

TPU continues to have no effect on U.S. imports from countries such as Taiwan in either subsample, as seen in columns 3 and 4. Moreover, in the last two columns we pool Taiwan and China and include HS-6 effects. This controls for the possibility that variation within high (or low) sunk cost industries have some unobserved correlation with the uncertainty measure. We continue to find a positive and significant differential effect of TPU on Chinese trade (relative to Taiwan) in high sunk cost industries. These estimates also include a full set of sector-country dummies with standard errors clustered on industry. Because the sunk cost measure is a generated regressor, we also computed standard errors using 500 bootstrap replications, which did not change the significance levels reported in this table.

In Panel B we split high and low sunk cost industries and analyze Chinese exports to the U.S. versus the EU. Recall we found no average effect of U.S. TPU on Chinese exports to the EU. We also find no differential effect of TPU on Chinese exports to the EU between high and low sunk costs (columns 3 and 4). We pool Chinese exports to the U.S. and EU in columns 5 and 6 and include sector-country and HS-6 industry dummies—thus controlling for industry specific shocks in China and any component of sunk costs common to both markets. We find a positive and significant differential effect on China's export growth to the U.S. in the high sunk cost industries and none in the low, as the model predicts.[34]

In sum, there is strong support for the sunk cost channel and the effect is robust to controlling for unobserved demand and supply shocks. This is the main source of heterogeneous TPU effects in our model and thus our focus. However, richer models could yield additional testable predictions, e.g. whether TPU is stronger in capital intensive industries.[35]

Timing of TPU shocks and pre-accession growth trends

We examined the effects of TPU on export growth between 2000-2005, which assumes a specific pre- and post-shock period. We examine the timing assumptions as follows. First, we ask if there are pre-existing trends, which would weaken the assertion that the shock was due to WTO accession. Second, we allow the data to identify when exactly the shock occurred.

Pre-accession growth trends could also generate an omitted variable bias if they persisted and were correlated with the uncertainty measure. To examine this we first run our baseline estimation on pre-accession Chinese trade. In Table A3, column 3 we find no significant effect of the uncertainty measure in 1996 on Chinese trade growth in 1996-99. To eliminate any HS-6 industry growth trends that persist from the pre-accession period we subtract the pre-accession equation (in changes) from the baseline 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.4. Columns 1 and 2 of Table A3 show the baseline results are not driven by pre-accession growth trends.

The results thus far focus on specific years and a balanced panel. For comparison to our earlier results we hold fixed the profit loss measure calculated using applied tariffs from 2000. In Online Appendix C.5 we show that the structural interpretation of the full panel estimates is $b_{\gamma t}^{panel} = \frac{k-\sigma+1}{\sigma-1}\frac{\beta g \lambda_2}{1-\beta}\Delta\gamma_t$, where $\Delta\gamma_t = \gamma_{2000} - \gamma_t$ for any year t. The estimates from Table A4 are plotted in Figure A1 and show no significant difference in the TPU effect in the pre-accession period: 1996-2001. This indicates that minor

[34]We employ the same sunk cost measure as in panel A, which requires only a similar industry ranking of these costs when exporting to the U.S or the E.U.

[35]The estimates in Table A5 provide some, albeit weak, evidence for this and are discussed in the appendix.

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.[36] Those beliefs seem to have been revised only after China accedes to the WTO. 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 Partial Effect Estimates of TPU on Prices and Entry

Prices

We now examine the effect of uncertainty on prices. The model predicts that a reduction in TPU lowers the ideal consumer import price index due to the entry of new varieties and a reduction in the prices of existing ones if there is technology upgrading. For industry V in state s this index is defined by $P_{sV,x} \equiv \left[\int_{\Omega^x_{sV}} (p_{vs})^{1-\sigma}\right]^{1/1-\sigma}$.[37] In Appendix B.4, we establish that changes in this index between 2005 and 2000 have an exact log linear representation in terms of uncertainty and trade costs, yielding

$$\Delta \ln P_{V,x} = -\left(\frac{k}{\sigma-1}-1\right)(-\ln U_V) + b^P_\tau \Delta \ln \tau_V + b^P_d \Delta \ln D_V + b^P + e_V. \tag{12}$$

where b^P contains the aggregate terms and e_V is an error term. Higher initial uncertainty, i.e. lower U, generates lower price changes because $k > \sigma - 1$ (for a finite first moment of exports). We use the linear approximation of $\ln U_V$ similarly to the OLS gravity estimates of eq. (11).[38]

We compute the exact change in the price indices for each industry using the approach in Feenstra (1994), which captures the contributions from new varieties and changes in the price of existing ones. We define varieties at the HS-10 product-country level, as in Broda and Weinstein (2006). We continue to assume a common σ so any correlation between the price change and our measure of TPU does not reflect variation in the elasticity.

Similarly to export growth, we first provide non-parametric and semi-parametric evidence. First, the distribution of price changes is lower for high TPU industries and present across the full range of TPU, as seen in figures 3(b) and 4(b). Second, in figure 4(d) we show the result is also present in the semi-parametric estimates after controlling for tariffs, transport costs and sector effects.

Table 7 presents the estimates for eq.(12). We find support for the prediction that industries with higher initial TPU had larger price reductions. In column 1 we see the uncertainty effect is significant when we define industries at the HS-6 level, as we do in the gravity equation. The number of observations is lower than the gravity regressions because quantity is not always consistently observed and because the price index is only defined for HS-6 industries where at least one variety is traded in both periods, which may introduce a sample selection bias. Any such bias should be mitigated by recomputing the price change at a higher

[36] These insignificant changes in γ during 1996-2001 are also consistent with the lack of variability in the vote share to revoke MFN status in the house of representatives. According to the Congressional Quarterly Almanac that share increased slightly from 33% in 1996 to 40% in 1997 and remained around that level (except for 2000, 34%). We also constructed and found that a news index of U.S. TPU did not fall significantly during 1996-2002 but did so between 2002-2006, both of which are consistent with the panel estimates of changes in γ. We thank two referees for these suggestions.

[37] These reflect the consumer prices inclusive of any tariffs and transport costs. A change in this price index could still be consistent with a negligible change in the aggregate U.S. price index if the expenditure share on those goods is negligible.

[38] Thus $\ln U_V \approx -\frac{ug}{\sigma-1}\left(1-\left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-3}\right)$ and the estimated OLS coefficient is $b^P_\gamma = -ug\frac{k-(\sigma-1)}{(\sigma-1)^2}$.

level of aggregation, such as the HS-4 level. In column 3 we do so and regress it on the HS-4 average of the TPU and trade cost measures. We find that the baseline results for uncertainty are robust to the level of aggregation. Moreover, the price effects are robust to controlling for sector effects (columns 2 and 4).[39]

The partial effects of uncertainty on prices are the following. First, using column 1, the average reduction in the price index is 15 log points at the mean of initial TPU. Second, the aggregate import price index change is a weighted average of the industry level changes; applying those same weights to the estimates of TPU in each industry we obtain its impact on the aggregate import price index: 17 log points. The price effects are stronger for the HS-4 regression, but we focus the quantification and robustness on the more conservative HS-6 estimates.

The baseline price index results are also robust to controlling for HS-6 industry-specific shocks. Examples of such shocks include changes in tastes or consumer demand for quality that could be correlated with initial uncertainty and thus bias our estimates. We apply the same approach used in Table 5. For each HS-6 we compute an additional import price index, which reflects all non-preferential partners trading with the U.S. and pool them with the Chinese observations to estimate the differential effect of TPU. The first column of Table A6 shows that industries where Chinese exporters faced higher initial TPU had larger price reductions if the good was imported from China but that same measure of TPU had no significant effect on the price index measure from the remaining countries. In column 2 we find a significant differential effect of TPU on the Chinese goods prices even after controlling for HS-6 industry effects. We find similar results in columns 3 and 4 when we restrict the control group to Taiwan. All the results control for sector-country growth shocks and cluster standard errors at the HS-6 level.

Evidence on Channels: entry, upgrading and sunk costs

The price index effects are the relevant ones for consumer welfare. But we can provide some additional evidence for the entry and upgrading channels and the role of sunk costs highlighted by the model.

The model predicts that at least a fraction $G(c_{sV}^U)$ of Chinese firms in industry V export to the U.S. In Appendix 2.2, we establish that changes in entry between 2005 and 2000 have an exact log linear representation in terms of uncertainty and trade costs, yielding the following estimation equation:

$$\Delta \ln n_V = k\left(-\ln U_V\right) + b_\tau^n \Delta \ln \tau_V + b_d^n \Delta \ln D_V + b + e_V \tag{13}$$

where b captures aggregate changes and e_V is an error term. The predicted signs match those for the export equation. Using the changes in the number of traded HS-10 products as a proxy for entry we find TPU reductions lead to entry in the same sample where we found it reduced the price index. Moreover, all other variables are significant and have the predicted sign (Table 8).

The magnitude of the uncertainty coefficient in column 1 of Table 8 is attenuated towards zero because entry is measured with error when we use product data. The attenuation reflects measurement error whenever a "true" variety is defined at a level finer than the HS-10, thus we do not use these estimates for quantification.

[39] To understand the potential bias note that in order for $\Delta \ln P_{V,x}$ to be defined for a given V at least one of its varieties must be traded in both periods. This is less likely at the narrower HS-6 industry level than the HS-4. Thus the HS-6 sample is less likely to include industries that are small and have higher variety churning. In a period of high entry this selection could attenuate the estimated impacts relative to the HS-4 sample. The stronger point estimates for uncertainty in column 3 suggest that is the case. The other potential concern with the HS-6 results is that they are more susceptible to measurement error and outlier influence since they average over fewer HS-10 price changes than the HS-4. We address this in columns 1 and 2 by trimming the top and bottom 2.5 percentile but obtain similar baseline results if we include these observations.

This issue is potentially important in industries where all HS-10 categories within an industry already had positive trade in both periods, so the growth in the number of measured variety is zero but true variety may have increased or decreased. We address this in column 3 by dropping those industries; the sign and significance are similar to those in the full sample but the impact of uncertainty triples. The results are robust to controlling for sector effects (column 4).

Directly testing the effect of TPU on upgrading requires data on export technology expenditures, which is not available. Nevertheless, we can provide indirect evidence. In the absence of upgrading the model predicts no effect of uncertainty on the prices of continuing varieties. The ideal price index for Chinese imports in industry V is composed of a weighted average of changes in unit values of HS-10 varieties traded in 2000 and 2005 (a measure of the change in average unit values) plus an adjustment term capturing the growth in the trade share of continuing varieties. If a reduction in TPU only operated by inducing export entry, then the model predicts an increase in the continuing variety component for industries with higher initial TPU, i.e. an increase in average prices since the entrants would be less productive.[40] In figure A2 we find the opposite: the local polynomial fit of the continuing variety component against TPU shows a significant negative relationship, which is also robust to controlling for changes in tariffs, transport costs and sector effects in a parametric setting similar to the one used for the full price index in Table 7.[41] This evidence that continuing varieties have substantially lower prices in industries with higher TPU is not consistent with a basic version of our model where TPU only affects entry; the finding requires a channel whereby reductions in TPU lower prices. One such channel is export technology upgrading.

In Table 9 we examine whether the baseline price and entry effects are stronger in high sunk cost industries. We interact the TPU measure with indicators for high and low sunk cost industries, defined as before. We find significant effects of TPU on prices and entry for high sunk cost industries but no significant effects for low, as the model predicts.

In sum, we have documented a strong and robust relationship of TPU on export values and prices of Chinese goods sold in the U.S. and provided evidence for the channels highlighted by the model. There is a non-negligible share of expenditure on those goods, which suggests this episode affected aggregate outcomes in the U.S. To quantify these we must model the general equilibrium effects of TPU.

4 General Equilibrium Effects of Policy Uncertainty

To quantify and decompose the GE effects of TPU we now allow the exporter to be large enough to affect aggregate outcomes in the destination market. We examine exports and domestic outcomes including consumer welfare and firm outcomes (entry investments, sales, employment). The exposition in the text focuses on the key equilibrium conditions, the expressions used for quantification, and the intuition for these in a single industry setting. The appendix provides additional details on the derivation of certain expressions and the extension to a multi-industry setting. We show that with limited information the model can be used to examine counterfactuals beyond the Chinese episode such as the impact of U.S. TPU against all its trade partners. In section 5 we employ the estimated structural parameters to examine the implications for China's WTO accession.

[40] We describe this further in the appendix (see equation (55)).

[41] This is true at the HS-4 or HS level or if we restrict ourselves to the uncensored entry sample in Table 8 (results available upon request).

4.1 Setup

The following additional assumptions allow us to determine aggregate expenditure, E_s, and price index, P_s, in a tractable way:

A1. There is no borrowing technology available across periods.
A2. Individuals are either workers, mass L, or entrepreneurs, mass N. Entrepreneurs are endowed with a blueprint embodied in the marginal cost, c_v, and receive the profits of their variety and any lump-sum rebates of tariff revenues.
A3. The period utility reflects a constant expenditure share on differentiated goods equal to $\mu > 0$ for workers and zero for entrepreneurs.
A4. There are two countries with identical preferences.

Under these assumptions TPU does not affect aggregate expenditures, which allows us to focus on the effects via the price index—the latter are important in our empirical setting and in understanding welfare effects for consumers. To see this clearly we highlight the following implications. (1) A1 implies that current expenditures must equal current income each period for each individual. (2) A2 implies that the only source of worker income is the wage, which is pinned down by the marginal product of labor in the numeraire—unity.[42] (3) The constant equilibrium wage and A1-A3 together imply that expenditure on differentiated goods is constant: $E_s = \mu L$ for all s; so the price index is the only aggregate endogenous variable that is uncertain in each country. (4) The indirect utility for workers is $\tilde{\mu} P_s^{-\mu}$ in each state.[43] (5) A3 implies that entrepreneurs have linear utility so the entry decision of risk neutral entrepreneurs is obtained by solving the Bellman equation defined in section 2.2.[44] (6) A4 rules out third country effects, mostly for expositional reasons.[45]

The price index for differentiated goods in state s depends on imported and home varieties, $\Omega_s = \Omega_{s,x} \cup \Omega_{s,h}$:

$$P_s^{1-\sigma} = \int_{\Omega_s} (p_{vs})^{1-\sigma} \, \mathrm{d}v = \int_{\Omega_{s,x}} (\tau_m d c_v/\rho)^{1-\sigma} \, \mathrm{d}v + \int_{\Omega_{s,h}} (c_v/\rho)^{1-\sigma} \, \mathrm{d}v \tag{14}$$

where $\rho \equiv \frac{\sigma-1}{\sigma}$. Before deciding to enter, firms form rational expectations about the expected price index, P_s^e. In equilibrium $P_s^e = P_s$ 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 $\mathbf{i}_s$, that includes: (i) the fixed exogenous parameters of the model including the survival rate and the time invariant set of potential varieties in each country, Ω; (ii) the structure of the model including the entry decision rules; (iii) the current realization of the policy, and; (iv) the equilibrium set of varieties sold in each market in the previous period, denoted by Ω_{t-1}. The state, s, is defined by the combination of the realized policy at t and Ω_{t-1}.

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, $\mathbf{q}_s$; (b) a market entry decision for each

[42] This follows because the population in each country is sufficiently large for the numeraire to be produced in equilibrium.

[43] The constant is $\tilde{\mu} \equiv w \ell \mu^{\mu} (1-\mu)^{(1-\mu)}$ where ℓ is the labor endowment and $w = 1$ in equilibrium.

[44] Namely, equation (4) evaluated at a, K, β for the export decision to the home market and the same equation evaluated at a_{sh}, K_h, β_h for domestic production entry. We can interpret the exogenous discount factor in the Bellman equation as the survival probability of the entrepreneur or more generally the product of that probability and the probability of survival of the invested entry capital. We rule out the possibility that entrepreneurs are credit constrained by assuming that their endowment ℓ is always at least as high as the sunk entry investment so they can always self-finance this cost in a single period even if it exceeds that period's operating profits.

[45] If third countries don't face tariff shocks themselves in this market then they can easily be included since any shocks to their competitors' tariffs affect third country firms only via the price index and so they react similarly to domestic firms.

differentiated firm v and a distribution of active firms, Ω_s, with prices, $\mathbf{p}_s$; (c) an expected and actual price index, P_s^e 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_s^e, the policy regime (Λ) and its lump-sum revenue, and all other information in $\mathbf{i}_s$; (iv) $P_s = P_s^e$ due to rational expectations (see Online Appendix A.2.2), and (v) the labor market clears.

Since TPU now affects the price index there will be aggregate uncertainty and transition dynamics. We derive key analytical results for cutoffs, to compare with the partial effects, and for the price index, to compare with the literature on the aggregate gains from trade. We then provide a numerical solution for the model using exact changes.

4.2 Equilibrium Entry, Prices and Welfare

The quantification computes exact changes, $\hat{y}_s \equiv y_s/y_b$— the ratio of some outcome y in state s and its baseline value. To fix ideas and simplify notation we choose a baseline with a deterministic policy $\tau = \tau_b$, so $y_b \equiv y\left(\tau_b, \gamma = 0\right)$. Using the demand expressions and eq. (14) we express the exact change in the aggregate price as a function of changes in two prices—one for imports, $P_{s,x}$, and the other for domestic varieties $P_{s,h}$ —where $P_{s,i} \equiv \left[\int_{\Omega_{s,i}} \left(p_{vs}\right)^{1-\sigma} dv\right]^{1/(1-\sigma)}$, $i = x, h$. Thus we have

$$\left(\hat{P}_s\right)^{1-\sigma} = I\left(\hat{P}_{s,x}\right)^{1-\sigma} + (1-I)\left(\hat{P}_{s,h}\right)^{1-\sigma} \quad \text{all } s \tag{15}$$

where the weight, $I \equiv \tau_b R_b/E_b$, is the share of expenditure on imported goods in an observed baseline period, i.e. the import penetration ratio, as shown in Appendix A.2.2.

Any firm with costs below the relevant cutoff in state s serves the market as do other surviving firms that previously entered under better conditions. Therefore the varieties that determine each $P_{s,i}$ can reflect the entry cutoffs in that state and prior ones. The expression in (15) reflects this hysteresis due to sunk costs as does the quantification. The latter focuses mainly on comparing stationary equilibria and so does the exposition below. Stationary equilibria are characterized by a constant mass of active domestic and foreign firms in each market and thus constant price indices for any given τ_m. The stationary price index is then $P_m = P\left(c_m, c_m^h, \tau_m\right)$, which evaluates (14) at the cutoffs for each state.[46] Similarly, the domestic component of the price index is $P_{m,h} = P_h\left(c_m^h\right)$ so its change, $\hat{P}_{m,h}$, depends on changes in the domestic entry cutoff, $\hat{c}_{m,h}$. The import price index is $P_{m,x} = P_x\left(c_m, \tau_m\right)$ and therefore $\hat{P}_{m,x}$ depends on the change in the cutoff and also in tariffs if $\tau_m \neq \tau_b$.

The key analytical results do not rely on a specific productivity distribution but the estimation and quantification uses a Pareto, and in this case we obtain

$$\left(\hat{P}_m\right)^{1-\sigma} = I\left(\hat{\tau}_m\right)^{1-\sigma}\left(\hat{c}_m\right)^{k-(\sigma-1)} + (1-I)\left(\hat{c}_{m,h}\right)^{k-(\sigma-1)} \quad \text{all } m \tag{16}$$

To determine the price and welfare effects we solve for the change in entry, first under the deterministic baseline, and then under uncertainty.

[46] If $T \geq 0$ periods ago the tariff changed to τ_m then the stationary equilibrium is given by the value of $y_{T\to\infty}\left(\tau_m\right)$. In a stationary equilibrium there is still exogenous death but it is exactly offset by entry thus leaving the firm mass unchanged.

Deterministic Policy Baseline

If the policy is expected to remain at τ_m then the price index is $P_m^D \equiv P\left(c_m^D, c_{m,h}^D, \tau_m\right)$, which evaluates (14) using the available varieties. An exporting firm serves this market if its cost is below $c^D\left(P_m^D, \tau_m\right)$, as given in eq.(3). A domestic firm serves the market if its cost is below $c_h^D\left(P_m^D\right) = \left[\frac{a_{mh}}{(1-\beta_h)K_h}\right]^{\frac{1}{\sigma-1}}$.

Each firm's investment to serve a country is independent of the price index in another country.[47] We can then solve for the price index, production and export cutoffs for each country separately by finding the unique solution to the three equations defining the equilibrium value of these variables in each country. We prove this in the appendix and show that higher tariffs increase the equilibrium price index, $dP^D/d\tau > 0$, reduce export entry, $dc^D/d\tau < 0$, and increase domestic entry, $dc_h^D/d\tau > 0$. Replacing the equilibrium cutoff changes in (16) we obtain the equilibrium change in the price index in any state m relative to a baseline state

$$\hat{P}_m^D = \left[I\left(\hat{\tau}_m\right)^{1-\frac{\sigma k}{\sigma-1}} + (1-I)\right]^{-1/k} \quad \text{all } m. \tag{17}$$

Consumer welfare in this model is simply $\left(\hat{P}_m^D\right)^{-\mu}$ and we can compare it to the cost of autarky, which is $\left[\hat{W}_m = \left(\hat{P}_m^D\right)^{-\mu}\right]_{\hat{\tau}_m\to\infty} = (1-I)^{\mu/k}$. Autarky is costlier, or equivalently the gains from trade with a given country larger, the higher the initial import share from that country, I, and the lower the trade cost elasticity, k. Thus the expression for deterministic welfare gains from trade for consumers in our model is similar to those obtained in a broader class of static trade models (cf. Arkolakis et al, 2012).[48]

Uncertain Policy

We now examine entry and prices when tariffs are expected to change with probability γ. Since we allow for three states, switching out of the intermediate state can worsen conditions for both foreign firms (to high protection) or domestic ones (to low protection). Either switch leads to gradual exit and transition dynamics as the price index adjusts up towards its stationary value. Therefore, firm decisions to enter in the intermediate state depend on the expected transition paths for economic conditions.

The total change in the exporter entry cutoff under uncertainty relative to the deterministic value under a baseline tariff τ_b is $\hat{c}_1 \equiv c_1^U/c_b^D$. It is derived by solving the optimal stopping problem in section 2.2 after allowing tariffs to also affect economic conditions indirectly through P. In Appendix A.2.3 we derive:

$$\hat{c}_1 = U\left(\omega g, \gamma\right) \times \hat{P}_1 \times \left(\hat{\tau}_1\right)^{-\frac{\sigma}{\sigma-1}} \tag{18}$$

$$U\left(\omega g, \gamma\right) \equiv \left[\frac{1+u\left(\gamma\right)\omega g}{1+u\left(\gamma\right)}\right]^{\frac{1}{\sigma-1}} \tag{19}$$

The effect of introducing uncertainty holding tariffs fixed is obtained by using $\hat{\tau}_1 = 1$. We note two differences relative to the small exporter case. First, there is a change in the price index, captured by $\hat{P}_1$. Second, the

[47] The separability arises because the equilibrium wage is constant and the marginal domestic entrant pays the sunk cost after knowing its productivity and is assumed to be unproductive enough that it never exports.

[48] Equation (17) also applies to a model with multiple exporters to this market if they face deterministic tariffs and have the same distribution parameter, k. We use I equal to either the aggregate import share (complete autarky) or the import share of a specific country (partial autarky).

average operating profits under the worst case scenario relative to state 1 is now $\omega g = (\tau_2/\tau_1)^{-\sigma} g$, where

$$g \equiv (1-\beta)\sum_{T=0}^{\infty}\beta^T \left(\frac{P_{2,T}}{P_1}\right)^{\sigma-1}, \tag{20}$$

and $P_{2,T}/P_1$ is the relative price index T periods after switching to high protection. So g is the average change in profits after a transition to high protection due to aggregate price changes. With a small exporter $g = 1$, whereas now it will typically differ from unity. Our solution will show that this GE effect does not overturn the direct tariff effect, i.e. we continue to have $(\tau_2/\tau_1)^{-\sigma} g < 1$ and thus $U < 1$. So, conditional on the price index change, we continue to obtain the TPU augmented gravity equation (10); the main difference is that the coefficient on τ_2/τ_1 reflects $u \times g$, whereas it reflected only u in the small exporter case.

Since $U < 1$, the direct effect of uncertainty is to reduce export entry, for given P. To determine the full export entry effect we must solve for $\hat{P}_1$, which in turn requires the change in domestic entry. With sunk costs of domestic entry, $K_h > 0$, the domestic cutoff change, $\hat{c}_{1,h} \equiv c^U_{1,h}/c^D_{b,h}$ is

$$\hat{c}_{1,h} = U_h(g_h,\gamma) \times \hat{P}_1 \tag{21}$$

$$U_h(g_h,\gamma) \equiv \left[\frac{1+u_h(\gamma)\,g_h}{1+u_h(\gamma)}\right]^{\frac{1}{\sigma-1}} \tag{22}$$

$$g_h \equiv (1-\beta_h)\sum_{T=0}^{\infty}(\beta_h)^T \left(\frac{P_{0,T}}{P_1}\right)^{\sigma-1}. \tag{23}$$

The relevant domestic TPU factor is now U_h, which depresses domestic entry. The intuition is similar to export entry, except that the worst case scenario for domestic firms is the low protection state. Starting in the intermediate state the expected duration of low protection is $u_h(\gamma) \equiv \gamma(1-\lambda_2)\frac{\beta_h}{1-\beta_h}$ and the expected change in profits after that transition, g_h. After a transition to low protection the domestic price index falls and thus so do domestic profits so $g_h < 1$. Figure 5 uses the parameters subsequently estimated and discussed and shows the transition paths for prices. The bottom line represents $P_{0,T}/P_1$ after a switch to the low protection state; the larger initial decline reflects the immediate entry of exporters (now facing lower protection) and the sluggish domestic exit (due to the hysteresis effect). Over time domestic firms die and fewer re-enter since the lower protection in the new stationary equilibrium will entail a lower price index. The top curve in figure 5 shows the transition after a switch to high protection; the undershooting in this case is due to the slow exit of foreign firms.

Using the entry cutoffs and (16) we obtain the change in the stationary price index between a baseline deterministic policy state and the intermediate one with TPU:

$$\hat{P}_1 = \left[I(\hat{\tau}_1)^{1-\frac{\sigma k}{\sigma-1}}(U)^{k-(\sigma-1)} + (1-I)(U_h)^{k-(\sigma-1)}\right]^{-1/k}. \tag{24}$$

We verify directly that when $U = U_h = 1$ we obtain the deterministic expression in (17). The price change due to TPU alone, if $\hat{\tau}_1 = 1$, increases the price index due to its effect on foreign firms, $U < 1$, and domestic ones, $U_h < 1$. Therefore the stationary intermediate state equilibrium under TPU must contain fewer foreign firms, fewer domestic firms, or both.

One may conjecture that TPU would be a form of protection that would promote domestic and hinder foreign entry, but this is not always the case. If the intermediate tariff value is very close to its low protection value, i.e. τ_1 is close to τ_0, then there are more domestic firms under TPU. This will be the case in our application to China and occurs because the domestic firms were already close to their worst case scenario of low protection. As a result, U_h is sufficiently close to unity and thus offset by the GE price effect in eq. (21). The price effect in this case is driven by the lower number of foreign firms, which fear a large reduction in profits when τ_1 is closer to τ_0 and relatively farther from τ_2. Conversely, if τ_1 were close to τ_2, then there would be more foreign entry under TPU and less domestic entry.

In sum, TPU increases the price index but its general equilibrium effects on entry are less obvious. The model predicts that reducing TPU can promote both foreign and domestic entry, as one of our counterfactual exercises will show. We will also ask if that reduction in domestic entry can ever be so large as to push consumer welfare under TPU below the permanent autarky level.

4.3 Solution

To fully characterize and quantify the effects of TPU we must solve the model for the complete sequence of prices and entry decisions. We describe the key elements and solution approach for the multi-industry version here and provide details in the appendix.

- **Inputs:** the model and its solution require
  - A set of exogenous parameters: $\Theta \equiv \{k, \sigma, \Lambda(\tau_m, \gamma), \beta, \beta_h\}$
  - Baseline equilibrium import shares: $\mathbf{I} \equiv \{I_V(\tau_b, \gamma = 0)\}$, where $I(\tau_b, \gamma = 0) = \Sigma_V I_V(\tau_b, \gamma = 0)$.
- **Equilibrium:** using the entry conditions in eqs. (18) and (21) and the definitions for U and U_h we obtain a non-linear system of equations for
  - the relative stationary price index in the intermediate state: $\hat{P}_1(g, g_h, \Theta, \mathbf{I})$ in eq. (24).
  - the sequence of relative prices after a switch to low or high protection, respectively $\hat{P}_{0,T}\left(g_h, \hat{P}_1, \Theta, \mathbf{I}\right)$, eq. (44) and $\hat{P}_{2,T}\left(g, \hat{P}_1, \Theta, \mathbf{I}\right)$, eq(45) in appendix A.2.2.
  - the average profit change due to prices after a switch to high or low protection, respectively $g\left(\hat{P}_{2,T}/\hat{P}_1, \Theta\right)$ in (20) and $g_h\left(\hat{P}_{0,T}/\hat{P}_1, \Theta\right)$ in (23).

    where $\hat{P}_{\cdot}$ denotes a price index relative to the baseline.
- **Solution**: $\Upsilon(\Theta, \mathbf{I}) \equiv \left\{\hat{P}_1; g; g_h; \left(\hat{P}_{2,T}; \hat{P}_{0,T}\right)_{T=0}^{\infty}\right\}$ found by
  - Fixing a set Θ consistent with our estimation and data $\mathbf{I}$.
  - Iterating n times until we obtain a fixed point such that $\Upsilon^{(n)}(\Theta, \mathbf{I}) = \Upsilon^{(n-1)}(\Theta, \mathbf{I})$.

To understand the approach recall that entry requires firms to incorporate expected changes in profits. Starting in state 1, a component of that change is captured by g and g_h. The initial values used in the solution algorithm are the upper bounds $g^{(0)} = \left(P_2^D/P_1^D\right)^{\sigma-1}$ and $g_h^{(0)} = \left(P_0^D/P_1^D\right)^{\sigma-1}$, which we compute using the deterministic equation in (16).[49] Using these we compute the initial values for $\hat{P}_1$ and the paths $\left(\hat{P}_{2,T}; \hat{P}_{0,T}\right)_{T=0}^{\infty}$. These sequences are then used to update g and g_h and we iterate until $g^{(n)} = g^{(n-1)}$ and $g_h^{(n)} = g_h^{(n-1)}$, at which point the prices also converge, so $\Upsilon(\Theta, \mathbf{I}) = \Upsilon^{(n)}(\Theta, \mathbf{I})$.

We use the following inputs. First, we use import expenditure shares in the baseline year, I_V, since these are the theoretically consistent industry weights for any policy terms in the price expressions. Second, we

[49] These are upper bounds because $P_1^D < P_1^U$ and because $P_{2,T}$ and $P_{0,T}$ converge respectively to P_2^D and P_0^D from below.

choose a post-agreement baseline year, 2005, for τ_{0V} and I_V. Third, we require values for Θ; the elasticities can be estimated (as we show below); β and β_h are obtained from annual export and domestic firm survival rates data. The tariffs for each of the three states are observable in the China episode (τ_{1V} is the 2000 MFN and τ_{2V} the column 2) and we will use these throughout but will also consider the effect of alternative counterfactual values.

The remaining key inputs to solve the model are the expected durations embodied in u and u_h. One contribution of this paper is to identify these parameters and quantify the role of TPU in China's WTO accession, which we do in section 5. But first we solve the model and describe its outcomes under a range of alternative uncertainty parameters and policies.

4.4 Outcomes and Policy Experiments

We solve the model under alternative policy regimes to explain its key mechanisms and implications. We also illustrate the model's applicability to alternative counterfactual policy experiments, including settings where some parameters have not yet been estimated. The main counterfactual we consider is one where the U.S. unilaterally abandons its trade agreement commitments against all its trading partners and imposes a policy regime similar to the one it used for China prior to 2002. This is a potentially useful benchmark to illustrate the qualitative effects of TPU and also to contrast its quantitative welfare effects with the costs of autarky, since the latter are well known and understood in this class of models in the absence of uncertainty.

To solve the model we fix all parameters other than the policy regime at the levels described in Table A9. We allow for alternative policy regimes as follows. Recall that $u \equiv \gamma\lambda_2 \frac{\beta}{1-\beta}$ and $u_h \equiv \gamma(1-\lambda_2)\frac{\beta_h}{1-\beta_h}$ so to pin down their values we require the probability of a policy change, γ, and the probability of high tariffs conditional on that change, λ_2. To focus on the uncertainty parameter we can fix a value for λ_2 and then solve the model under all possible $\gamma \in [0,1]$. In section 5, when we apply the model to China we use (i) $\hat{\tau}_{1V}$ and $\hat{\tau}_{2V}$ as the ratio of the 2000 MFN and column 2 tariffs respectively to the 2005 MFN tariffs in each industry V, and; (ii) Chinese import penetration in each industry V. In this section we examine the effect of introducing TPU on all U.S. trade partners. We assume that τ_{0V}, τ_{1V} and τ_{2V} are identical across all V and set them equal to their simple means when computing $\hat{\tau}_1$ and $\hat{\tau}_2$. Therefore we only require aggregate import penetration.

For any given set of parameters, the qualitative impacts of TPU towards all trade partners are similar to those we later obtained for China alone. But since aggregate import penetration is large, 0.26 in 2005, introducing TPU on all partners generates stronger aggregate impacts on U.S. prices and firm outcomes. Our objective is to explore the implications of the model under alternative policy regimes. We focus on the range of possible outcomes rather than any specific value arising from the quantification.

In this section we address the following counterfactual: what would be the effects of the U.S. abandoning its trade agreements in 2005, raising its applied tariff slightly, and introducing the possibility that it would either start a trade war or return to the agreement state. In this section we assume high and low protection are equally likely, setting $\lambda_2 = 1/2$, and vary the probability of a policy change, γ. In contrast, in section 5 we use specific parameters implied by the estimation for China. We proceed in three steps.

1. Evaluate the effects of switching policy states within a regime, e.g. transition to the low protection state, and decomposing them into a TPU and an applied policy component.
2. Perform counterfactual analysis of switching policy regimes, e.g. the effects of introducing TPU under

different applied or threat policies.

3. Calculate gains from trade under TPU and draw implications for trade agreements.

Effects of switching policy states

The total change in the stationary price index between state 1 and 0, where we interpret the latter as an agreement, can be decomposed into a TPU and applied policy change as follows:

$$\hat{P}_1 \equiv \frac{P_1}{P_0^D} = \underbrace{\left(\frac{P_1}{P_1^D}\right)}_{\text{TPU}} \times \underbrace{\left(\frac{P_1^D}{P_0^D}\right)}_{\text{Applied } \hat{\tau}_1}, \tag{25}$$

To obtain the TPU component we solve the model for $\hat{P}_1$, at each possible γ, and then divide it by the applied policy component, $\frac{P_1^D}{P_0^D} = \hat{P}_1^D$ computed from (16). We can decompose the impact of TPU on entry and all other variables similarly. The observed tariffs in 2005 are very similar to the ones in 2000 so $\hat{\tau}_1$ and thus $\hat{P}_1^D$ are very small, so nearly all of the effect of the agreement is from TPU changes. Thus, in figure 7 we focus on the TPU components of variables, which we graph against all possible γ.

- **Aggregate price index:** increases by as much as 5% if $\gamma = 1$ with almost half that increase occurring even at moderate uncertainty ($\gamma = .25$).
- **Foreign sales and entry:** fall by as much as 35% (sales) and 60% (entry) if $\gamma = 1$. Even moderate uncertainty generates a considerable reduction in entry, 40%.
- **Domestic sales and entry:** sales increase by up to 12% if $\gamma = 1$. Entry increases due to TPU but there is an inverse U-shape. At sufficiently high γ the direct effect of TPU, U_h, starts to offset the indirect price effect. Below we show that if applied tariffs in the intermediate case were not so close to those in the low state then the direct effect dominates.

We break the price index change into its foreign and domestic components in Figure 6. The decline in foreign variety entry due to TPU causes the foreign component to increase by as much as 30%. This large change is partly offset in the aggregate price index due to the decline in the price index of domestic varieties.

For any particular value of γ we can also plot the transition path for the price index relative to state 1 as we do in figure 5 using $\gamma = 0.25$. The dashed line represents the price index after a switch to the higher protection state and the solid bottom line to the low protection state.

Effects of switching policy regimes

Different events can trigger a change in γ without any change in a state or tariff values. For example, in the years leading up to WTO accession Chinese exporters may have changed their assessment about the probability of that outcome. Alternatively, if the U.S. abandoned its trade agreements we could consider the impact of changes in γ. The graphs just described also allow us to evaluate such counterfactuals by simply taking the ratio of the outcomes at different γ since $\hat{P}_1(\gamma')/\hat{P}_1(\gamma) = P_1(\gamma')/P_1(\gamma)$.

An alternative counterfactual regime is one with different threat tariffs. To understand its effects we continue to fix $\tau_1 = 1.04$, we also fix γ and then compute $\hat{P}_1$ and $\hat{P}_1^D$ by solving the model at all alternative counterfactual values of τ_2. We obtain the TPU component for prices using (25) and do the same for the other outcomes. The first column in figure 8 contains the results. We use $\gamma = .25$ and allow $\tau_2 \in [1.04, 1.38]$

so it ranges from τ_1 to the value of τ_2 in the data that were used in figure 7.[50] Thus the outcomes at the maximum threat in Figure 8 are, by construction, exactly the same as those in Figure 7 when $\gamma = .25$.

In Figure 8 we verify that TPU has the strongest effects on entry and sales at the highest threat tariff. Reductions in τ_2 have opposing effects on domestic and foreign entry and thus on their respective price indices. But the import price index effect always dominates; which we can see in Figure 8(a) where the aggregate price effect of TPU is monotonic in τ_2. At the highest τ_2 introducing TPU implies an increase in the price index that is about 1/3 of what would result if the U.S. reverted to autarky. The autarky cost is shown by the straight line and is computed using eq. (17).

Finally, we consider the impact of TPU under alternative applied tariffs and address two issues. First, if the U.S. did abandon its trade commitments it could revert to a temporary tariff that is higher than its MFN in 2000; we show that in this case TPU can reduce both foreign and domestic entry. Second, we isolate a pure risk effect of TPU.

We fix τ_2 and γ and solve the model for $\tau_1 \in [\tau_0, \tau_2]$. The results in the second column of Figure 8 reflect two effects of increasing τ_1. First, at higher τ_1 the import penetration on which TPU acts is lower. Second, higher τ_1 implies a relative decrease in the threat for foreign varieties and the opposite for domestic varieties. We summarize the effects of TPU at different τ_1 as follows.

- **Aggregate price index:** increases due to TPU at all τ_1, but slightly less so at higher τ_1 because import penetration is lower.
- **Foreign sales and entry:** both fall with TPU when τ_1 is close to τ_0 (as seen before) since the possibility of high protection implies a substantial tariff increase. This negative effect of TPU is reversed when τ_1 approaches τ_2. At that point foreign exporters have little to lose if the policy switches to τ_2 and thus the direct price effect (due to lower entry of domestic firms) dominates.
- **Domestic sales and entry:** both increase with TPU when τ_1 is close to τ_0 (as seen before). But if τ_1 is above 1.1 then TPU reduces entry because the direct effect, from lower U_h, eventually offsets the indirect price effect. At high enough τ_1, TPU reduces domestic entry and can even reduce domestic sale values.

One of the counterfactuals uses a value of τ_1 that is equal to the long-run mean of the policy, which is useful to isolate the pure risk effect of TPU. In this exercise the long-run mean is $\bar{\tau}_1 = \lambda_2 \tau_2 + (1 - \lambda_2)\,\tau_0 = 1.21$ so introducing TPU at that point can be interpreted as a pure risk effect, since it is a mean preserving spread of the policy. The outcomes in Figure 8 evaluated at that mean show that the pure risk effect of TPU is to lower both foreign and domestic entry.

In the quantification section for China we also show that the counterfactual at the mean tariff can be used to determine what fraction of TPU effects can be attributed to a pure risk effect.

Gains from trade, value of agreements and tariff bindings

What are the implications of these results for the value of trade agreements and some of their key features such as tariff bindings?

An immediate implication is that to the extent that agreements reduce TPU then they reduce domestic prices and increase consumer welfare. How important are the aggregate price effects of TPU relative to say

[50] The qualitative results are similar if we use alternative interior values of γ. At $\gamma = .25$ the probability of high protection, $\gamma\lambda_2$, is similar to what we subsequently estimate for China.

imposing prohibitive tariffs? This depends on the policy regime. At the baseline value of τ_1 used in figure 6 autarky generates a price increase of about 6.8%; in figure 7(a) we see the effect of TPU is 1/3 of that when $\gamma = .25$ and 2/3 if $\gamma = 1$. This suggests an important value of agreements that eliminate such uncertainty.[51]

Can the cost of TPU ever exceed that of autarky? In figure 8(b) we see that it may, depending on the initial tariff. The price effect of autarky, $\left(\hat{P}_m^D\right)_{\hat{\tau}_m \to \infty} = (1 - I_{\tau_1})^{-1/k}$, is decreasing in τ_1 because at higher tariffs there is lower import penetration and thus a lower cost of eliminating trade. At high enough τ_1 the cost of TPU is higher than that of autarky because TPU reduces domestic entry by so much that it eventually leads to less entry than autarky.

The possibility that TPU is costlier than autarky relies on TPU reducing domestic entry. In this exercise that occurs above τ_1=1.1. During most of the GATT era the U.S. simple average tariff has been below 1.1, but it was around 1.22 immediately preceding GATT, in 1947 (Bown and Irwin, 2015). So if GATT 1947 reduced the probability of a trade war, which was one of its objectives, then the model suggests it may have increased both foreign and domestic entry investments and realized a large fraction of the possible gains from trade (since at $\tau_1 = 1.22$ the price effect of TPU is close to that of autarky).

If we re-interpret the model then our counterfactual results for τ_2 also provide support for the emphasis the WTO places on negotiating reductions in tariff bindings. In the WTO countries commit not exceed bound tariff rates, but have discretion to set their applied tariffs anywhere below them. We can re-interpret the model as corresponding to three different states between members of the WTO: in state 1 countries have discretion to set tariffs anywhere at or below the binding, τ_2, and a probability $\gamma\lambda_2$ they will use the discretion and set τ_2; state 0 corresponds to giving up any such discretion. After the Uruguay Round the U.S. applied and tariff bindings are almost the same, so this would match the outcome when $\tau_2 = \tau_1$ in Figure 8. Under this interpretation the results show that uncertainty shocks have stronger effects for countries with higher bindings and thus increasing the binding at a given fixed tariff decreases trade and welfare. Equivalently, they show that past negotiations to reduce bindings alone can increase trade substantially, as found by Handley (2014) for Australia.

In sum, the policy experiments in this section illustrate how the model works qualitatively and how it can be applied more broadly. The range of outcomes we obtain indicate an important role for trade agreements. To narrow the range of outcomes we now turn to a specific episode where we estimate the uncertainty parameters.

5 Structural Estimates and Quantification

NLLS Structural Estimates

We identify the key structural parameters to quantify the effects of TPU via non-linear estimation. This approach differs from section 3 in two ways. First, while the export equation is still given by (10), the uncertainty factor, $U(\omega_V g, \gamma)$, now reflects a general equilibrium factor common to all industries. Second, we now use $f(.) = -(k - \sigma + 1)\ln U(\omega_V g, \gamma)$ and the definition of U to rewrite eq. (11) in terms of estimable

[51] In our setting the welfare effects for U.S. consumers would be similar if other countries retaliated and introduced uncertainty on U.S. exporters. The latter would have lower profits. This is due to the separability of markets, fixed wage and the assumption that the marginal domestic entrant knows its productivity and is not an exporter.

coefficients instead of using a linear approximation around no uncertainty. We obtain

$$\Delta \ln R_V = b_{d\sigma} \ln \left[\frac{1 + \tilde{b}_\gamma \left(\frac{\tau_{2V}}{\tau_{1V}} \right)^{-b_\sigma}}{1 + \tilde{b}_\gamma / g} \right] + b_\tau \Delta \ln \tau_V + b_d \Delta \ln D_V + b + e_V \tag{26}$$

where the parameters have the following structural interpretation: $b_d = -k$ and $b_\tau = -\frac{k\sigma}{\sigma-1}$, $\tilde{b}_\gamma = u(\gamma) g$, $b_\sigma = \sigma$, and $b_{d\sigma} = -\frac{k-\sigma+1}{\sigma-1}$. One component of U, $1 + \tilde{b}_\gamma / g = 1 + u$, is log-separable and does not vary by industry so we cannot identify it separately from the constant, b. The non-linear baseline regression estimates $\tilde{b}_\gamma$, b_d and b and imposes two theoretical restrictions: $b_{d\sigma} = \frac{(b_d + b_\sigma - 1)}{b_\sigma - 1}$ and $b_\tau = b_d \frac{b_\sigma}{b_\sigma - 1}$; as before we impose $b_\sigma = 3$ but we will now test it.

Column 1 of Table 10 provides non-linear least squares (NLLS) estimates. For comparison with earlier results we transform the estimate for $\tilde{b}_\gamma = 0.736$ into its OLS regression counterpart, $b_\gamma^{nlls} = \frac{-b_d - b_\sigma + 1}{b_\sigma - 1} \tilde{b}_\gamma = 0.90$. This is slightly higher than its the OLS estimate (column 2) and significantly different from zero. Results are similar when we control for sector effects (columns 3 and 4).[52]

Consistency with model and other evidence

Before using these estimates we ask whether they are consistent with the model and other evidence. The signs of all estimated parameters are those predicted by the model. We re-ran the NLLS baseline specifications by individually relaxing $b_\tau = b_d \frac{\sigma}{\sigma-1}$ or $b_\sigma = 3$. We fail to reject either of these restrictions and report p-values in the last two rows of Table 9. 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 estimates that use similarly disaggregated U.S. trade and tariff data.[53]

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 $[1.4, 3.1]$. The estimate is larger than 1 and it satisfies the model's requirement for a finite first moment of sales, which we did not impose. The magnitude is similar to what is found by other studies using firm level data.[54]

The other parameter central to the quantification is u—a Chinese exporter's expected duration of a spell under column 2 prior to WTO accession. We can obtain a lower bound for $u = \tilde{b}_\gamma / \bar{g}$, without numerically solving the model by computing the upper bound $\bar{g} = \left(P_2^D / P_1^D\right)^{\sigma-1}$, based on the deterministic price change formula in (16). We find this lower bound to be very similar to the estimate, $\hat{u} = \tilde{b}_\gamma / \hat{g} = .73$, where $\hat{g} = 1.004$ from solving the model with the estimated parameters as described in the previous section. This expected duration estimate reflects the exporters' beliefs for an event that never took place. We can't defend a particular value. Nevertheless, the bounds on our estimate are reasonable and consistent with the model. To fix ideas, consider a Chinese firm that starts exporting in 2000. Since firm level studies suggest an expected export duration of between 6 and 7 years, our estimate implies that those firms expected to spend at least

[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 U.S. statutory tariffs and HS-6 imports.

[54]Eaton et al (2011) obtain an aggregate estimate of 2.46 using French exports; Hsieh and Ossa (2015) obtain a range from 1 to 1.44 over industries using Chinese firm data.

10% of their exporting spell under column 2.[55]

We also estimate the probability of transitioning from state 1 to 2, given by $\lambda_{12} = \gamma\lambda_2$. Using the definition of $\tilde{b}_\gamma$ and u we obtain the estimate $\hat{\lambda}_{12} = \frac{1-\beta}{\beta}\hat{u}$. The estimation does not impose any restriction on this parameter, but we find that it is bounded in the unit interval for reasonable annual probabilities of firm survival, β. The point estimate using the value for β employed in the general equilibrium derivation, is $\hat{\lambda}_{12} = 0.13$.[56]

In the presence of endogenous domestic firm entry we require those firms' beliefs of the expected duration of an agreement, u_h, to compute general equilibrium effects. We are unable to identify this parameter empirically using U.S. firm entry or sales data because the relevant uncertainty factor, U_h, does not vary across industries (all the tariff effects work through the price index). Therefore we parametrize u_h by defining $\alpha \equiv u_h/u$. We choose $\alpha = 4$ as the central value, which implies that before the agreement a U.S. firm expected to spend 4 times as long under the WTO state than a Chinese exporter expected to spend under column 2. In section 5.1 we show the quantification results are not very sensitive to alternative feasible values of α.[57]

We also solved a special case of the model without domestic entry costs, so there is a constant mass of active domestic producers, which is independent of u_h. We find similar export value effects, and the effects for the aggregate price index are only somewhat stronger than under endogenous domestic entry (because the absence of a domestic extensive margin is partially offset by a larger intensive margin impact).

A final cross-validation of the NLLS estimates is to ask if the implied uncertainty measure, $\hat{U}_V$, can predict the observed industry price index changes exactly as predicted by eq.(12). In section 3.6 we estimated eq.(12) and found larger price reductions in industries with higher initial TPU by using a linear approximation to U_V. Our objective here is to test a more specific structural implication of the model. We regress the observed changes in prices on tariff and trade cost changes and on $\left(\frac{k}{\sigma-1} - 1\right)(-\ln U_V)$, where the latter is constructed using the NLLS estimates obtained using the export data. The uncertainty measure thus constructed is predicted to have a coefficient of -1 on price changes and we estimate it to be -0.96 (s.e. 0.21). Thus the estimates from the export equation predict the effect of TPU on prices quite well.

5.1 Quantification: TPU and Exports

We quantify the effects of TPU changes, decompose them into a pure risk and mean effect, perform different counterfactual experiments and provide an ad valorem tariff equivalent cost of TPU. Table A9 summarizes the parameters we use based on our estimates and auxiliary data.

Similarly to section 4.4 we quantify the effect of re-introducing uncertainty in 2005, but now on Chinese exporters only. Qualitatively, the outcomes for Chinese varieties are similar to to section 4.4 but there are two differences worth noting. First, export entry and sales reflect the response of Chinese varieties whereas the "domestic" entry and sales reflect both U.S. and other non-Chinese varieties.[58] Second, we must now

[55] The export survival is 6.25 years under an exit rate of 0.16 (the fraction of new Chinese exporters that stop exporting after one year as reported by Ma et al., 2014) and 6.7 if $1 - \beta = 0.15$ (the value we use in the quantification).

[56] When $\hat{u} = .73$ then $\hat{\lambda}_{12} = .73\frac{1-\beta}{\beta} \in (0, 1]$ if $\beta \in [.42, 1)$.

[57] Our estimate of $\gamma\lambda_2 = 0.13$ implies that $\lambda_2 \in [0.13, 1]$ and so, after applying the discount factor values, the range consistent with the estimates is $\alpha \in [0, 12]$, for which we report sensitivity. For the central case, $\alpha = 4$ we obtain $\lambda_2 = 0.28$.

[58] The U.S. firms respond to the general equilibrium price index changes, as they did in the two-country model, but so do any other non-Chinese firms that face constant trade barriers in the U.S. It is simple to show that, because we are solving for changes, this is the outcome from extending the model to multiple countries as long as all non-Chinese firms face a common σ

account for variation in policy across industries and aggregate in a theoretically consistent way. The main quantitative difference is that in section 4.4 we did not have the uncertainty parameters and so focused on describing a range for the effects of TPU. We can now use our estimates for China to pinpoint a particular value for each outcome. We now contrast these GE point estimates to the partial effect estimates and place them into perspective relative to the observed changes during this period.

In the discussion that follows we quantify the effect of re-introducing TPU in 2005, e.g. R_1/R_1^D for exports, and interpret its inverse as the impact of TPU reduction. We present all results using log changes.

Average Effects: General vs. Partial

The average log change from increasing TPU is given by

$$\mathbb{E}_V \ln R_{1V}/R_{1V}^D = \mathbb{E}_V \ln (U_V)^{k-(\sigma-1)} + k \ln P_1/P_1^D. \tag{27}$$

The direct effect is -30 log points and price index effect is 2 log points. Therefore the model implies that the reduction in TPU in 2000-2005 lead to an average increase of 28 log points, as shown in Table 11. [59]

To highlight the quantitative importance of accounting for non-linearities and price effects we contrast it with the partial effect from the linear estimation. The latter assumes no price effects and a linear approximation to U so the estimated partial effect counterpart to eq. (27) is simply $-b_\gamma^{OLS} \times \mathbb{E}_V \left(1 - \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-3}\right) = -36$ log points. Thus the implied partial effect in exports due to a TPU reduction is more than 20% higher than the GE effect. All subsequent exercises focus on the GE non-linear estimates.

Aggregate Effects

We now turn to aggregate effects of TPU for exports. The direct uncertainty effect, $-\ln (U_V)^{k-(\sigma-1)}$, ranges from about 0 to 57 log points in the data and there is also considerable dispersion in export shares, so the simple average growth can underestimate the aggregate effect if uncertainty is higher in relatively larger industries. Thus we need to appropriately weight each uncertainty factor by the relevant expenditure share to compute the growth of total expenditure on Chinese goods due to TPU[60]

$$\ln R_1/R_1^D = \ln \sum_V r_{1V}^D (U_V)^{k-(\sigma-1)} + k \ln P_1/P_1^D. \tag{28}$$

This aggregate effect is 32 log points, slightly higher than the average effect in Table 11.

The quantification implies that TPU can account for over a third of observed changes on expenditure in Chinese goods. The counterfactual holds income and aggregate U.S. expenditure on differentiated goods constant, so it also applies to the growth in the share of U.S. tradeables expenditure on Chinese goods, i.e. the growth in Chinese import penetration. In Table 11 we report this share increased by 73 log points so TPU can account for over a third of that growth.

and k, as is typical in this type of model, and a similar expectation of the duration of the low protection state, u_h.

[59] We compute $(U_V)^{k-(\sigma-1)}$ using the estimates in column 1 of Table 10, the implied $\hat{u} = .73$, and each τ_{1V}/τ_{2V}. The price effect term uses the multi-industry version of (24) at fixed tariffs and the decomposition given in (25). The effect is not sensitive to alternative values of α, it ranges from 29 to 30 log points if $\alpha = 6$ or 0 respectively.

[60] The appropriate weight to evaluate the TPU component is the import share at τ_1 that would have been observed if $\gamma = 0$, $r_{1V}^D \equiv \tau_{1V} R_{1V}^D / \Sigma \tau_{1V} R_{1V}^D$. Since these are not observed we compute them by using the observed initial equilibrium imports, R_{0V}^D, the model implied deterministic change due to tariff changes, $\hat{\tau}_{1V}$, and the estimated tariff elasticity, b_τ. Thus we obtain $r_{1V}^D = \tau_{1V} R_{0V}^D (\hat{\tau}_{1V})^{b_\tau} / \Sigma \tau_{1V} R_{0V}^D (\hat{\tau}_{1V})^{b_\tau}$ since the aggregate effects cancel out.

We noted that, for any given level of aggregate expenditure, the effect of TPU on import penetration growth is the same as its effect on imports. So introducing TPU in 2005 would reduce penetration by 32 lp from 4.5% to 3.3%. The effect of TPU can depend on the baseline year's expenditure share, which may vary for exogenous reasons. Thus we calculate the effect for import penetration for each year after the agreement until 2010 and plot them in Figure 1. The solid line shows the observed import penetration, which tripled between 2000 and 2010. In contrast, the dashed line is the GE counterfactual showing instead that it would only have doubled if TPU had remained.

Mean-risk Decomposition

In section 4.3 we describe how to decompose the effect of an agreement into a change in applied policies and TPU. The latter TPU effect can be further decomposed into changes in the mean vs changes in risk. To show this, we rewrite the impact of a change in γ evaluated at the pre-agreement applied tariff level, τ_{1V} as follows

$$\ln \frac{R_V(\gamma_0, \tau_{1V})}{R_V(\gamma_1, \tau_{1V})} = \ln \frac{R_V(\gamma_0, \bar{\tau}_V)}{R_V(\gamma_1, \bar{\tau}_V)} + \left[\ln \frac{R_V(\gamma_0, \tau_{1V})}{R_V(\gamma_0, \bar{\tau}_V)} - \ln \frac{R_V(\gamma_1, \tau_{1V})}{R_V(\gamma_1, \bar{\tau}_V)} \right] \text{ each } V. \tag{29}$$

The first term on the right is the growth in exports due to credibly securing tariffs at their long-run expected value, denoted by $\bar{\tau}_V$, a reduction in γ will then work as a mean preserving compression in tariffs and thus we label this the risk effect. If $\tau_{1V} < \bar{\tau}_V$ then eliminating uncertainty reduces the expected value of tariffs and this lock-in effect is captured by the term in brackets, which is positive when the initial tariffs are below the long-run mean, as is the case in our application.

To quantify the risk component we require the counterfactual long-run tariffs, $\bar{\tau}_V$. Prior to the agreement these are $\bar{\tau}_V = \lambda_2 \tau_{2V} + (1 - \lambda_2)\, \tau_{0V}$ and can be computed using an estimate of λ_2 and the observed values for τ_{2V} and τ_{0V}.[61] In section 4.3 we showed that the TPU effect can be computed at any counterfactual τ_1. We set $\tau_1 = \bar{\tau}_V$ and compute $\ln \frac{R_V(\gamma_0, \bar{\tau}_V)}{R_V(\gamma_1, \bar{\tau}_V)}$ by industry. We then aggregate these changes, using the expenditure weights at $\bar{\tau}_V$. This yields a risk component of TPU for exports equal to 23 log points, which is 71 percent of the total. The substantial share due to risk arises because even if we start at the higher mean tariffs the threat of moving to τ_2 entails a doubling of tariffs on average.[62]

Ad valorem equivalents: TPU vs. applied policies

To compare the effects of TPU with other policies we calculate ad valorem tariff equivalents (AVE) of TPU on exports and other outcomes. The AVE is defined as the deterministic log change in the uniform tariff factor, $\ln \Delta_y$, that generates the same change in an outcome y as TPU. Formally, Δ_y is the implicit solution to

$$y(\tau_1 \Delta_y, \gamma = 0) = y(\tau_1, \gamma > 0)\,. \tag{30}$$

If we divide both sides by the baseline value for exports $R(\tau_1, \gamma = 0)$ then the expression on the RHS will yield the 32 log point change due to TPU that we previously derived. The LHS will then reflect the change in exports due to a deterministic tariff change, both the direct and indirect effect via prices. Solving for the AVE we obtain $\ln \Delta_R = 5$ log points (Table 12). So the export AVE was higher than the U.S. applied average tariff in 2000, which was about 4 log points, and also higher than the AVE of U.S. applied non-tariff barriers on manufacturing (e.g. anti-dumping, licensing, etc) as calculated by Nicita et al. (2009).

[61] When the extreme states are absorbing the long-run mean is equal to the mean conditional on exiting MFN.

[62] Our baseline uses $\alpha = 4$, which implies $\lambda_2 = .28$, but we also find large risk shares for alternative values that are consistent with our empirical estimates, e.g. for an α of 2 or 6 ($\lambda_2 = .44$ or $.21$), the risk shares are between 54 and 78 percent .

5.2 Quantification: Entry, Prices and Welfare

We now quantify the aggregate effects of TPU on entry, U.S. price indices, and additional domestic outcomes. We then compare their magnitude with other pieces of evidence.

Aggregate Price and Welfare

The overall price index increase from re-introducing TPU in 2005, $\ln P_1/P_1^D$, is 0.52 log points.[63] In Table 12 we show this is equivalent to a deterministic tariff increase of 13 log points using (17) and the AVE definition in (30). This is roughly half the price index increase predicted by the model if there was no TPU and the U.S. stopped all imports from China in 2005.

Moreover, increasing TPU at the (counterfactual) mean tariff increases the price index by 0.515 log points—almost the same as increasing it at the lower MFN tariffs. This implies the lock-in effect defined in equation (29 from eliminating TPU on the price index is close to zero because it generates an increase in foreign varieties that is offset by a decrease domestic varieties.[64] Thus the pure risk cost of TPU on the price index is large.

These price index effects also apply to the (stationary) effect of TPU on consumer welfare because it is simply $-\mu \ln P_1/P_1^D$. Namely, the welfare cost of TPU in 2005 is almost half that of going to autarky with China. The price index AVE also applies to consumer welfare. Thus the effect of a TPU increase on the price index and welfare is equivalent to permanently raising average tariff protection in 2005 from 4 to about 17 log points on Chinese goods.

The AVE and autarky comparisons for welfare are relative magnitudes and so independent of the U.S. expenditure share on manufacturing, μ. To provide an absolute effect and place it in context of other large trade shocks we use the 2005 U.S. expenditure on manufacturing as a share of tradeables expenditure and obtain a welfare cost of TPU of 0.45 log points.[65] This is over half the cost that Costinot and Rodríguez-Clare (2014) calculate for North America under a worldwide trade war with a uniform tariff of 40%.[66] The TPU effect is also substantial when compared to another reference point for the magnitude of gains from trade: Broda and Weinstein (2006) estimate the real income gain from all new imported varieties in the U.S. between 1990-2001 to be 0.8 percent.

Other outcomes

In our setting the price AVE also applies to various outcomes of incumbent U.S. firms in the differentiated sector, namely their domestic profits, sales, and employment, which are affected by uncertainty and tariffs only indirectly via P. The aggregate effects of TPU on domestic firm outcomes also reflect changes in entry decisions. The latter depend directly on uncertainty so their AVE is different from the one for the price index, as shown in Table 12. Re-introducing TPU in 2005 would increase U.S. firm entry by 0.44 log

[63] We compute it using the multi-industry version of (24) at fixed tariffs and the decomposition given in (25).

[64] This does not mean that changes in tariffs have little effect on the price index. In fact, the price impact of increasing tariffs to their mean in the absence of uncertainty is 0.43 log points. However, if we lower those tariffs back down under uncertainty the price index will fall by almost as much so the net lock-in effect is close to zero.

[65] Over 98% of Chinese exports to the U.S. are manufactures. As is standard in most trade models neither our quantification nor the ones discussed below take services into account. 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 relatively low values of U.S. imports/consumption captured by the import penetration.

[66] They find it is 0.7 percent; our models differ in some dimensions: e.g. uncertainty, sunk costs and an outside good.

points—equivalent to a 1.8 log point increase in tariffs. The effect on aggregate domestic sales is about 1.3, equivalent to a tariff increase of 5.9. Finally, TPU increases domestic quantities and employment in the differentiated sector by 1.2 log points, equivalent to a 7.3 log point tariff increase on Chinese goods, a sizable permanent tariff change in the context of recent U.S. agreements.

How do these outcomes relate to observed changes in 2000-2005? To answer this we identify the differentiated sector with manufacturing, and the numeraire with the remaining tradeable sectors (agriculture and mining), as is standard. In this period, we observed a reduction in both gross and net entry of manufacturing establishments. According to the Business Dynamics Statistics Database (U.S. Census) the manufacturing gross entry rate was 7.9 percent and gross exit was 8.7 on average between 2002-2006.[67] We also observed an expansion in the non-manufacturing sector, both in terms of establishments and employment. The sign of these observed changes is predicted by the model after a reduction in TPU.

The quantification can also account for a non-trivial fraction of U.S. manufacturing employment and domestic sales reallocation. To see this recall that our counterfactuals hold total employment, L, constant. Thus the model predicts the reduction in TPU in 2000-2005 reduced the manufacturing employment *share* in the tradeable sector by 1.2 log points (Table 12). This share, which controls for any changes in employment in tradeables, fell by 3.3 log points in the data during this period. The U.S. sales counterfactual holds total manufacturing expenditure, $E = \mu w L$, constant and implies a 1.3 log point reduction in the U.S. firms' share of manufacturing expenditure; the observed reduction in that share was 3.3 log points. Thus the reduction in TPU can account for at least a third of the reallocation of domestic manufacturing sales and a similar fraction of its employment share in tradeables.[68]

Export Entry and Price Index

We conclude the section by comparing the quantitative implications for export entry and price indices with other sources of information.

The model predicts that at least a fraction $G(c_{sV}^U)$ of all Chinese firms in V export to the U.S. in state s. The growth in the export cutoff due to the agreement is in eq. (18) and its TPU component is obtained by holding tariffs fixed. Therefore, under a Pareto distribution, the growth in the number of exporters is simply k times that expression. Similarly to the average export effect in (27) we can compute the average entry effect of TPU as

$$\mathbb{E}_V \ln n_{1V}/n_{1V}^D = k\left(\mathbb{E}_V \ln U_V + \ln P_1/P_1^D\right). \tag{31}$$

On average TPU reduced entry by 54 log points. This is also the increase in the number of firms that upgrade since the fraction of exporters that upgrade is independent of uncertainty in this setting.

The aggregate entry and upgrading effect is 61 log points, which is sizable relative to the growth in the number of Chinese firms exporting to the world over 2000-2005 (83 log points, Ma et al., 2014).[69] The AVE for entry is nearly twice as large as for exports (Table 12). Moreover, when we apply the decomposition in eq. (29) to entry we find that most of the TPU effect on entry is attributable to a risk reduction.

[67]The magnitude of the decline in the actual number of establishments is not directly comparable to the counterfactual quantification because the latter keeps the mass of potential manufacturing firms constant whereas in the U.S. it is not.

[68]A more complete analysis of the effects of TPU on U.S. firm entry and employment requires extending the model to account for features such as the input-output linkages analyzed by Acemoglu et al. (2014).

[69]The aggregate entry effect recomputes the expression in (28) using the entry elasticity for U. The expenditure weights used are the relevant ones to obtain the effect of changes in entry on prices. If we had the number of exporting firms in each V in the initial period we could compute the growth in the total number of exporting firms due to TPU.

The negative effects of TPU on entry and upgrading imply increases in the price index for Chinese varieties. Comparing (15) and (16), the change in the import index for each industry when holding tariffs fixed is $\hat{P}_{1V,x} = (\hat{c}_{1V})^{1-\frac{k}{\sigma-1}}$. Thus the average effect of TPU across industries is

$$\mathbb{E}_V \ln P_{1V,x}/P^D_{1V,x} = \left(1 - \frac{k}{\sigma - 1}\right)\left(\mathbb{E}_V \ln U_V + \ln P_1/P^D_1\right), \tag{32}$$

which is simply a rescaling of the entry effect in (31). Using the NLLS estimates obtained with export data we compute the terms on the RHS and find an average price effect of 15 log points. The aggregate price effect is 17 log points and is obtained by aggregating the computed industry effects using the theoretically consistent weights in (28). In Table 11 we show that both the average and aggregate effects computed here are similar to the partial effect estimates we obtained in section 3.6 when using the empirical counterpart to this price index.

In sum, the model predicts that the reduction in TPU reduced the price index for Chinese varieties by 17 log points and has a small impact on non-Chinese varieties. This predicted relative price reduction is of a similar magnitude to the observed changes in this period. To see this we use the data in section 3.6 and compute the change in the ideal price of Chinese relative to non-Chinese imported varieties between 2000 and 2005. In Table 11 we show this is -15.4 log points.[70]

6 Conclusion

We assess the impact of trade policy uncertainty in a tractable general equilibrium framework with heterogeneous firms. 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 China's WTO accession and use it to estimate and quantify the impacts of reducing the trade policy uncertainty faced by Chinese exporters when the U.S. ended its annual threat to revert to Smoot-Hawley tariffs.

We derive observable, theory-consistent measures of TPU and estimate its effect on trade flows, prices and welfare. We find a large and robust effect of reducing TPU on China's export growth to the U.S. The same measure of TPU does not predict China's exports to other major industrial countries or U.S. import growth from other non-preferential U.S. trade partners. We also find that the reduction in TPU lowered Chinese industry export price indices, as the model predicts. Consistent with our model, these Chinese export value and price effects in the US market are strongest in industries with high sunk costs of exporting.

Using the estimates of the structural parameters we compute the exact changes in price indices and the effect on entry and sales of domestic and foreign firms. Had the MFN status been revoked, the typical Chinese exporter would have faced an average tariff of 31%. The removal of this threat had large effects on Chinese export entry, about 60 log points, and export growth, 32 log points or about 1/3 of the observed change. The quantification indicates the reduction in TPU decreased U.S. manufacturing sales and employment by more than one percent, but also lowered the price index and thus improved consumer welfare by the equivalent of a permanent tariff decrease of 13 percentage points on Chinese goods. Thus TPU had provided a substantial amount of effective protection, especially relative to the average applied tariffs, which were only about 4 percent.

[70] We obtain this by aggregating the industry price index changes for i =China or U.S. non-preferential trading partners using their respective log change ideal weights, $\Delta \ln P_i \equiv \sum_V w_{V,i} \Delta \ln P_{V,i}$. We obtain $\Delta \ln P_{china} - \Delta \ln P_{non-pref} = -15.6 + .2$.

Our findings have implications beyond this particular important event. They also indicate that an important role of agreements is to reduce policy uncertainty, which can be a substantial source of welfare gains, even if applied tariffs are unchanged. For example, we show that for a range of applied tariffs an increase in policy uncertainty may leave consumers worse off than autarky. It would be interesting for future work to explicitly model, quantify and decompose the relative importance of alternative channels through which TPU may operate, e.g. intermediates and offshoring; as well as its impact relative to that of alternative sources of Chinese export growth, e.g. changes in own trade policy, dismantling of central planning. It could also be useful to structurally quantify the labor market effects of TPU in the presence of frictions. More generally, our research points to the value of specific data-rich setting to identify the effects of policy uncertainty on economic activity and shows these potentially substantial effects should not be ignored.

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A Theory Appendix

A.1 Entry under Partial Equilibrium

This appendix derives the equilibrium entry expressions and results described in section 2.4, which are summarized in proposition 1 below. We derive these using the more general policy transition matrix below, and note the special case in the text is obtained when $\lambda_{22} = 1$.

$$M = \begin{bmatrix} \lambda_{22} & \lambda_{21} & 0 \\ \lambda_{12} & \lambda_{11} = 1-\gamma & \lambda_{10} \\ 0 & 0 & 1 \end{bmatrix} \tag{33}$$

Proposition 1: Policy Uncertainty and Export Entry (small exporter).
Under a regime $\Lambda(\tau_m, \gamma)$ with policy uncertainty and where tariff increases are possible, $\tau_2 > \tau_1$ and $u(\gamma) > 0$, the entry cutoff in the intermediate state, c_1^U, is
(a) unique and $c_1^U = c_1^D U(\omega, \gamma)$, and U is given by eq. (7)
(b) lower than the deterministic, $c_1^U < c_1^D$, and decreasing in policy uncertainty ($dc_1^U/d\gamma = dU/d\gamma < 0$ all γ)
(c) lower than the cutoff in the low state, $c_1^U < c_0^U = c_0^D$ and $c_1^U/c_0^U = c_1^U/c_0^D = U(\omega, \gamma) \times (\tau_1/\tau_0)^{-\frac{\sigma}{\sigma-1}}$.

Proof

To prove the uniqueness of the industry cutoff in (*1a*), we first establish sufficient conditions for a unique tariff below which each firm enters. We say $\Lambda(\tau_m, \gamma)$ exhibits uncertainty persistence if $\Lambda(\tau_{m+1}, \gamma)$ first order stochastically dominates $\Lambda(\tau_m, \gamma)$ for $m = 0, 1$, which is satisfied by (33).

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, $\tau_s^U(\gamma, c)$, below which a firm enters into exporting.*

Proof of Lemma 1*:* 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

$$\Pi(a_s, c, \gamma) - \Pi_e(a_s, c, \gamma) + K = \max\{0, \beta \mathbb{E}_s [\Pi(a_s', c, \gamma) - \Pi_e(a_s', c, \gamma)] - \pi(a_s, c) + K\} \tag{34}$$

$$V_s = \max\{0, \beta \mathbb{E}_s V_s' - \pi(a_s, c) + K(1-\beta)\} \tag{35}$$

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 $a_s \equiv EP^{\sigma-1}\tau_s{}^{-\sigma}\sigma^{-\sigma}(\sigma-1)^{\sigma-1}$ so $\mathbb{E}_s V_s' = \int V_s d\Lambda(\gamma, \tau'|\tau)$.

(2) Because $-\pi(a_s, c)$ is increasing in τ it is more attractive to wait at higher tariffs because the second element of (35) and therefore V_s would be higher, all else equal.

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

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

Proof of Prop. 1(a)

Lemma 1 shows that each firm v has a single tariff entry cutoff $\tau_s^U(\gamma, c_v)$. All firms have different cost but face the same τ and γ in the industry so there is a unique entry cutoff for any given τ_m, $c_1^U(\tau_m, \gamma)$, and only those with cost below this enter into exporting.

To show $c_1^U = c_1^D U(\omega, \gamma)$ we first derive $\Pi(a_s, c \leq c_s^U, \gamma)$ if $s = 1$. Starting with (35) and taking the expectation over the possible states we have

$$\begin{aligned}\mathbb{E}_s V_s' &= \lambda_{s,s+1}\left[\beta\mathbb{E}_{s+1}V_s' - \pi(a_{s+1}, c) + K(1-\beta)\right] \quad \text{if } c \leq c_s^U \\ &= \lambda_{s,s+1}\left[\beta\left(\frac{\lambda_{s+1,s+1}}{1-\beta\lambda_{s+1,s+1}}\left[K(1-\beta) - \pi(a_{s+1}, c)\right]\right) - \pi(a_{s+1}, c) + K(1-\beta)\right] \\ &= \frac{\lambda_{s,s+1}}{1-\beta\lambda_{s+1,s+1}}\left[K(1-\beta) - \pi(a_{s+1}, c)\right]\end{aligned} \tag{36}$$

where the second line uses (35) and takes the conditional expectation starting at $s+1$:

$$\begin{aligned}\mathbb{E}_{s+1}V_s' &= \lambda_{s+1,s+1}\left[\beta\mathbb{E}_{s+1}V_s' - \pi(a_{s+1}, c) + K(1-\beta)\right] \quad \text{if } c \leq c_s^U \\ &= \frac{\lambda_{s+1,s+1}}{1-\beta\lambda_{s+1,s+1}}\left[K(1-\beta) - \pi(a_{s+1}, c)\right]\end{aligned} \tag{37}$$

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

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

$$\pi\left(a_s, c_s^U\right) + \beta\lambda_{s,s+1}\frac{\pi\left(a_{s+1}, c_s^U\right)}{1-\beta\lambda_{s+1,s+1}} = K(1-\beta)\left(1 + \frac{\beta\lambda_{s,s+1}}{1-\beta\lambda_{s+1,s+1}}\right) \tag{38}$$

Starting at $s = 1$, replacing π with its value in (1) and simplifying we obtain the cutoff expression (6) in the text

$$\begin{aligned}a_1\left(c_1^U\right)^{1-\sigma}\left[1 + \frac{\beta\lambda_{12}}{1-\beta\lambda_{22}}\frac{a_2}{a_1}\right] &= K(1-\beta)\left(1 + \frac{\beta\lambda_{12}}{1-\beta\lambda_{22}}\right) \\ a_1\left(c_1^U\right)^{1-\sigma}\left[1 + u(\gamma)\frac{a_2}{a_1}\right] &= K(1-\beta)(1 + u(\gamma)) \\ c_1^U = \left(\frac{a_1}{K(1-\beta)}\right)^{\frac{1}{\sigma-1}} &\times \left(\frac{1+u(\gamma)\omega}{1+u(\gamma)}\right)^{\frac{1}{\sigma-1}} = c_1^D \times U_1(\omega, \gamma)\end{aligned} \tag{39}$$

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

Proof of Prop 1(b)

Since $c_1^U/c_1^D = U$ we must show $U < 1$ iff tariff increases are possible. From the definition in (7) we obtain $U < 1$ iff $u(\gamma)\omega < u(\gamma)$, which is true iff $\tau_2 > \tau_1$ and $\gamma\lambda_2 > 0$ so that $u(\gamma) > 0$.

Since $c_1^U = c_1^D U(\omega, \gamma)$ (part a) we can use (7) to obtain

$$\gamma \frac{d \ln c_1^U}{d\gamma} = \gamma \frac{d \ln U_1(\omega, \gamma)}{d\gamma} = \frac{1}{\sigma - 1} \frac{u}{1+u} \frac{\omega - 1}{1 + u\omega} < 0 \tag{40}$$

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

Proof of Prop. 1(c)

If $\tau_2 > \tau_1$ *and* $u > 0$ then $c_1^U < c_1^D$, (part b). Since $\lambda_{00} = 1$ 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 $c_1^U < c_1^D \leq c_0^D = c_0^U$. Using c_1^U from part (a) $c_1^U/c_0^D = U(\omega, \gamma) \times \left(c_1^D/c_0^D\right) = U(\omega, \gamma) \times (\tau_1/\tau_0)^{-\frac{\sigma}{\sigma-1}}$, where the last equality uses (3), and the definition of a_s for fixed E and P. **QED**

A.2 Entry and Prices under General Equilibrium

A.2.1 Derivation and comparative statics under deterministic policy baseline

The equilibrium baseline price index change in equation (17) and the comparative statics can be derived as follows. First, the price index is $P^D\left(c_m^D, c_{mh}^D, \tau_m\right) = \left[N \int_0^{c_m^D} (\tau_m d c_v/\rho)^{1-\sigma} dG(c) + N_h \int_0^{c_{mh}^D} (c_v/\rho)^{1-\sigma} dG_h(c)\right]^{1/(1-\sigma)}$. Second, we use the cutoff expressions, eq. (3) for exports and the counterpart evaluated at a_h, K_h, β_h for domestic firms. We can then write $c_{mh}^D = c_m^D \tau_m^{\frac{\sigma}{\sigma-1}} \left[\frac{(1-\beta)K}{(1-\beta_h)K^h}\right]^{\frac{1}{\sigma-1}}$, and reduce the system to two equations and show their unique intersection. For any given fixed tariff value the entry schedule, c_m^D, is linear and increasing in P_m^D and $c_m^D|_{P_m \to 0} = 0$ whereas $P^D\left(c_m^D, c_m^D \tau_m^{\frac{\sigma}{\sigma-1}} \left[\frac{(1-\beta)K}{(1-\beta_h)K^h}\right]^{\frac{1}{\sigma-1}}, \tau_m\right)$ is positive and decreasing in c_m^D. We replace each cutoff change in (16) to simplify to obtain (17).

A.2.2 Price index expectations, transition dynamics and exact changes

Expectations of future price index: P_s^e

Firms can derive P_s^e as follows. To predict the import component of P_s firms use the observed policy realization, τ_m, and must infer the set of exported varieties, Ω_s^x, over which to integrate. The latter is simply $\Omega_s^x = \Omega_s^{cont} \cup \Omega_s^{entry}$ where Ω_s^{cont} represents the set of foreign producers that exported to this market both in the previous and current periods (so $\Omega_s^{cont} = \emptyset$ in the initial trading period). The measure of continuers is given by the measure of previous period exporters—observed in Ω_{t-1}—adjusted by the exogenous survival probability, β, applied to all subsets. So Ω_s^{cont} is independent of the current tariff and economic conditions. New exporters are represented by the subset Ω_s^{entry} of all potential firms in the foreign country that (i) did not export in the previous period—known from Ω_{t-1}—and (ii) have a cost such that entry is optimal in state s according to (4). To predict the domestic component they do the same using the optimal cutoff obtained by solving the Bellman equation for the domestic entrant, given by (4) when evaluated at K_h, $a_{s,h}$ and β_h.

Transition Dynamics

Starting from the stationary equilibrium of the intermediate state 1 with cutoffs c_1^U and $c_{1,h}^U$, the price index for all $T \geq 0$ after switching to policy state $m = 0$ or $m = 2$ is

$$(P_{m,T})^{1-\sigma} = N\tau_m^{1-\sigma}\left(\int_0^{c_{m,T}} (c_v/\rho)^{1-\sigma}\, dG(c) + \beta^{T+1}\int_{\min\left\{c_{m,T},c_1^U\right\}}^{c_1^U} (c_v/\rho)^{1-\sigma}\, dG(c)\right) \tag{41}$$

$$+ N_h\left(\int_0^{c_{m,T,h}} (c_v/\rho)^{1-\sigma}\, dG(c) + \beta_h^{T+1}\int_{\min\left\{c_{m,T,h},c_{1,h}^U\right\}}^{c_{1,h}^U} (c_v/\rho)^{1-\sigma}\, dG(c)\right)$$

where in equilibrium we find $\min\left\{c_{m,T}, c_1^U\right\} = c_{m,T}$ if $m = 2$ (conditions worsen for foreign firms under high protection) and c_1^U otherwise and $\min\left\{c_{m,T,h}, c_{1,h}^U\right\} = c_{1,h}^U$ if $m = 2$ and $c_{m,T,h}$ otherwise. The representation holds for all $T \geq 0$ when states $m = 0, 2$ are absorbing.

Exact changes

Aggregate price index change and price sub-indices eq. (15).

We use the definition of P_s and rewrite it using the sub-indices $P_{s,i} \equiv \left[\int_{\Omega_{s,i}} (p_{vs})^{1-\sigma} dv\right]^{1/(1-\sigma)}$, $i = x, h$ and $\hat{y}_s \equiv y_s/y_b$

$$\begin{aligned}
(P_s)^{1-\sigma} &= (P_{s,x})^{1-\sigma} + (P_{s,h})^{1-\sigma} \\
\left(\frac{P_s}{P_b}\right)^{1-\sigma} &= \left(\frac{P_{s,x}}{P_b}\right)^{1-\sigma} + \left(\frac{P_{s,h}}{P_b}\right)^{1-\sigma} \\
\left(\hat{P}_s\right)^{1-\sigma} &= \left(\frac{P_{b,x}}{P_b}\right)^{1-\sigma}\left(\hat{P}_{s,x}\right)^{1-\sigma} + \left(\frac{P_{b,h}}{P_b}\right)^{1-\sigma}\left(\hat{P}_{s,h}\right)^{1-\sigma}
\end{aligned}$$

Eq. (15) follows once we recognize that $I_b \equiv \frac{\tau_b R_b}{E} = \left(\frac{P_{b,x}}{P_b}\right)^{1-\sigma}$. This equality is obtained from rewriting aggregate expenditure on imports and using the optimal demand in a baseline period:

$$\tau_b R_b = \int_{\Omega_{b,x}} p_v q_v = \frac{E}{P_b^{1-\sigma}}\int_{\Omega_{b,x}} p_v^{1-\sigma} \Rightarrow \frac{\tau_b R_b}{E} = \left(\frac{P_{b,x}}{P_b}\right)^{1-\sigma}$$

Stationary aggregate price index change as a function of cutoffs, eq. (16)

Above we show eq. (15) holds for all s so, under an unbounded Pareto distribution, eq. (16) holds for all *stationary* policy states m iff $\left(\hat{P}_{m,x}\right)^{1-\sigma} = (\hat{\tau}_m)^{1-\sigma}(\hat{c}_m)^{k-(\sigma-1)}$ and $\left(\hat{P}_{s,h}\right)^{1-\sigma} = (\hat{c}_{m,h})^{k-(\sigma-1)}$. For the foreign index we have

$$\left(\hat{P}_{m,x}\right)^{1-\sigma} = \frac{\int_{\Omega_{m,x}} (p_{vm})^{1-\sigma} dv}{\int_{\Omega_{b,x}} (p_{vb})^{1-\sigma} dv} = (\hat{\tau}_m)^{1-\sigma}\frac{\int_0^{c_m} c_v^{1-\sigma} dG(c)}{\int_0^{c_b} c_v^{1-\sigma} dG(c)} = (\hat{\tau}_m)^{1-\sigma}(\hat{c}_m)^{k-(\sigma-1)} \tag{42}$$

where the first equality is the definition, the second follows from replacing the optimal price and uses a constant cutoff due to the stationary equilibrium. The last equality uses the Pareto. Similarly we find

$$\hat{P}_{m,h} = (\hat{c}_{m,h})^{k-(\sigma-1)} \tag{43}$$

Deterministic price index change, eq.(17)

Substituting the deterministic cutoff from eq.(3) and the definition of a_m and doing similarly for an analogous expression for the domestic cutoff we obtain.

$$\begin{aligned}
\hat{c}_m^D &= (\hat{a}_m)^{\frac{1}{\sigma-1}} = (\hat{\tau}_m)^{\frac{-\sigma}{\sigma-1}}\hat{P}_m \\
\hat{c}_{m,h}^D &= (\hat{a}_{m,h})^{\frac{1}{\sigma-1}} = \hat{P}_m
\end{aligned}$$

replacing these in (16) and solving for $\hat{P}_m$ we obtain eq.(17).

General aggregate price index transition as a function of cutoffs

To derive an expression for P_{mT}/P_1 as a function of the cutoffs we derive $P_{m,T,i}/P_{1,i}$ and replace in eq. (15) to obtain

$$\left(\frac{P_{m,T}}{P_1}\right)^{1-\sigma} = I_1\left(\frac{P_{m,T,x}}{P_{1,x}}\right)^{1-\sigma} + (1-I_1)\left(\frac{P_{m,T,h}}{P_{1,h}}\right)^{1-\sigma}$$

Using the transition expression in (41) we can write

$$\begin{aligned}\left(\frac{P_{m,T,x}}{P_{1,x}}\right)^{1-\sigma} &= \left(\frac{\tau_m}{\tau_1}\right)^{1-\sigma}\frac{\int_0^{c_{m,T}} c_v^{1-\sigma}dG(c) + \beta^{T+1}\int_{\min\left\{c_{m,T},c_1^U\right\}}^{c_1^U} c_v^{1-\sigma}dG(c)}{\int_0^{c_1^U} c_v^{1-\sigma}dG(c)}\\ &= \begin{cases}\left(\frac{\tau_2}{\tau_1}\right)^{1-\sigma}\left(\left(1-\beta^{T+1}\right)\left(\frac{c_{2,T}}{c_1^U}\right)^{k-(\sigma-1)} + \beta^{T+1}\right) & \text{if } m=2\\ \left(\frac{\tau_0}{\tau_1}\right)^{1-\sigma}\left(\frac{c_{0,T}}{c_1^U}\right)^{k-(\sigma-1)} \quad \text{if } m=0\end{cases}\end{aligned}$$

$$\begin{aligned}\left(\frac{P_{m,T,h}}{P_{1,h}}\right)^{1-\sigma} &= \frac{\int_0^{c_{m,T,h}} c_v^{1-\sigma}dG(c) + \beta_h^{T+1}\int_{\min\left\{c_{m,T,h},c_1^U\right\}}^{c_{1,h}^U} c_v^{1-\sigma}dG(c)}{\int_0^{c_{1,h}^U} c_v^{1-\sigma}dG(c)}\\ &= \begin{cases}\left(\frac{c_{2,T,h}}{c_{1,h}^U}\right)^{k-(\sigma-1)} \quad \text{if } m=2\\ \left(1-\beta_h^{T+1}\right)\left(\frac{c_{0,T,h}}{c_{1,h}^U}\right)^{k-(\sigma-1)} + \beta_h^{T+1} \text{ if } m=0\end{cases}\end{aligned}$$

We use the stationary value of state 0 as a baseline, i.e. I_0, so below we rewrite $\hat{y}_s \equiv y_s/y_0^D$

$$\begin{aligned}\left(\frac{P_{m,T}}{P_0^D}\right)^{1-\sigma} &= I_0\left(\frac{P_{m,T,x}}{P_{0,x}}\right)^{1-\sigma} + (1-I_0)\left(\frac{P_{m,T,h}}{P_{0,h}}\right)^{1-\sigma}\\ \left(\hat{P}_{m,T}\right)^{1-\sigma} &= I_0\left(\frac{P_{m,T,x}}{P_{1,x}}\hat{P}_{1,x}\right)^{1-\sigma} + (1-I_0)\left(\frac{P_{m,T,h}}{P_{1,h}}\hat{P}_{1,h}\right)^{1-\sigma}\end{aligned}$$

Replacing eqs. (42), (43) and $P_{m,T,i}/P_{1,i}$ derived above and simplifying we have

$$\left(\hat{P}_{0,T}\right)^{1-\sigma} = I_0\left(\hat{c}_{0,T}\right)^{k-(\sigma-1)} + (1-I_0)\left(\left(1-\beta_h^{T+1}\right)\left(\hat{c}_{0,T,h}\right)^{k-(\sigma-1)} + \beta_h^{T+1}\left(\hat{c}_{1,h}\right)^{k-(\sigma-1)}\right) \tag{44}$$

$$\left(\hat{P}_{2,T}\right)^{1-\sigma} = I_0\left(\hat{\tau}_2\right)^{1-\sigma}\left(\left(1-\beta^{T+1}\right)\left(\hat{c}_{2,T}\right)^{k-(\sigma-1)} + \beta^{T+1}\left(\hat{c}_1\right)^{k-(\sigma-1)}\right) + (1-I_0)\left(\hat{c}_{2,T,h}\right)^{k-(\sigma-1)} \tag{45}$$

Multi-industry version

As we show below the domestic cutoff changes are function of aggregate variables. So the multi-industry version requires aggregation of only the foreign variables. We can then re-derive all the expressions by defining $\hat{P}_{s,x,V}$ at the industry level and aggregating the effects as required by the theory using the import share across industries: $r_{Vb} \equiv \tau_{Vb}R_{Vb}/\sum_V \tau_{Vb}R_{Vb}$.

$$\left(\hat{P}_s\right)^{1-\sigma} = I_b\sum_V r_{Vb}\left(\hat{P}_{s,x,V}\right)^{1-\sigma} + (1-I_b)\left(\hat{P}_{s,h}\right)^{1-\sigma}$$

Similarly for all other price expressions we replace the foreign variety variables such as cutoff changes by their mean using r_{Vb} as the weight.

A.2.3 Entry cutoffs

We derive the export and domestic cutoffs in the intermediate state presented in eqs. (18) and (21). We also derive their counterparts after a transition to either high ($m = 2$) or low protection ($m = 0$), which are used in the solution algorithm to obtain expressions for the transition prices in eqs. (44) and (45).

We focus on the comparisons of the steady state under intermediate protection with uncertainty ($m = 1$) versus without. Similarly to the partial effect derivation there is a positive probability of policy change at $m = 1$. The key difference is that now the exporter is large so after any change the domestic price index is affected and the exogenous death of firms generates transition dynamics. Thus the relevant states are no longer only $m = 0, 1, 2$. They are now $s = 1; m, T$ for $m = 0, 2$ and all $T \geq 0$ where T is the number of periods since the change from $m = 1$.

Transition cutoffs: $\hat{c}_{m,T}$

If $m = 0, 2$ are absorbing states then the sequence of business conditions, a_s, is deterministic for any $s = m, T$ and its path is determined by $P_{m,T}$ in eq. (41). Moreover, along the transition path the conditions are improving due to gradual exit (from exporters if $m = 2$ or domestic if $m = 0$) so $a_{m,T+1} > a_{m,T}$. Since conditions are improving but firms still face a risk of death they still have an option value of waiting. Therefore the marginal firm is the one indifferent between entering today and tomorrow so the future profit terms cancel and we obtain

$$\pi\left(a_s, c_s^U\right) / (1-\beta) = K \Leftrightarrow c_s^U = \left[a_s / (1-\beta) K\right]^{\frac{1}{\sigma-1}} \quad \text{if } s = 0, T; 2, T;$$

which has a similar functional form as the deterministic cutoff evaluated at current conditions.[71]

A similar expression applies to the cutoff for domestic firms: $c_s^U = \left[a_{s,h} / (1-\beta_h) K_h\right]^{\frac{1}{\sigma-1}}$. So we can rewrite either relative to some respective baseline and obtain $\hat{c}_s = \hat{a}_s^{\frac{1}{\sigma-1}}, \hat{c}_{s,h} = \hat{a}_{s,h}^{\frac{1}{\sigma-1}}$.

Intermediate state cutoff: exporter, $\hat{c}_1$

To obtain the formula for $\hat{c}_1 \equiv c_1^U / c_b^D$ in (18) we derive

$$c_1^U = \left[\frac{1 + u(\gamma)\,\omega g}{1 + u(\gamma)}\right]^{\frac{1}{\sigma-1}} \left[\frac{a_1}{(1-\beta) K}\right]^{\frac{1}{\sigma-1}} \tag{46}$$

and combine it with the definitions for a_1, c_b^D in eq.(3), g in eq.(20) and U in eq. (19). The derivation is identical to part (a) of proposition 1 except now we change (37) to reflect the transition dynamics in P after the tariff increases, so we have

$$\mathbb{E}_{s=2,T} V_s' = \lambda_{22}\left[\beta \mathbb{E}_{s=2,T+1} V_s' - \pi\left(a_{s=2,T+1}, c\right) + K(1-\beta)\right] \quad \text{if } c \leq c_s^U \tag{47}$$

Solving forward we obtain $\mathbb{E}_{s=2,0} V_s' = -\lambda_{22} \sum_{t=0}^{\infty} (\beta\lambda_{22})^t \pi\left(a_{s=2,T+1}, c\right) + \frac{\lambda_{22}}{1-\beta\lambda_{22}} K(1-\beta)$. Replacing this in (36), using the absorbing state, $\lambda_{22} = 1$, and simplifying we obtain

$$\mathbb{E}_{s=1} V_s' = \frac{\lambda_{12}}{1-\beta}\left[K(1-\beta) - (1-\beta)\sum_{t=0}^{\infty} (\beta)^t \pi\left(a_{s=2,T}, c\right)\right] \tag{48}$$

The cutoff expression for the marginal firm in $s = 1$ solves $V_1\left(c_1^U\right) = 0$, which we obtain as in proposition 1 but using (48):

$$\beta \mathbb{E}_{s=1} V_s'\left(c_1^U\right) - \pi\left(a_1, c_1^U\right) + K(1-\beta) = 0$$

[71] We prove this formally in the working paper for $m = 2$ with exogenous domestic entry. When domestic entry is endogenous then the initial price jump in the price index after a tariff increase is smaller but there is still gradual exit of exporters.

Using the definition of π, u and re-arranging we have

$$a_1\left(c_1^U\right)^{1-\sigma}\left[1+u(\gamma)(1-\beta)\sum_{t=0}^{\infty}\beta^t\left(\frac{a_{s=2,T}}{a_1}\right)\right]=K(1-\beta)(1+u(\gamma))$$

where the key difference relative to (39) is the term in [ ], which reflects average profits during the transition (instead of the fixed profits $\pi(a_2,c)$). Re-arranging and using the definitions of ω and g we obtain eq. (46).

Intermediate state cutoff: domestic, $\hat{c}_{1,h}$

The general entry problem for domestic firms is similar to the one for exporters (see section 2). The cutoff expression will differ in two ways. First, the domestic firms fear the low protection state rather than the high. Second, the deterioration in conditions for the domestic firms reflects only the general equilibrium effects due to entry of foreign firms and consequent reductions in the price index (it does not reflect a direct tariff effect).

To derive the cutoff we first write the option value of waiting for each potential domestic entrant (the domestic entry version of equation (35)):

$$V_s^h=\max\left\{0,\beta_h\mathbb{E}_sV_s^{h\prime}-\pi(a_{s,h},c)+K_h(1-\beta_h)\right\}$$

where $V_s^h\equiv\Pi(a_{s,h},c,\gamma)-\Pi_e(a_{s,h},c,\gamma)+K_h$ and $\mathbb{E}_sV_s^{h\prime}\equiv\mathbb{E}_s\left[\Pi(a'_{s,h},c,\gamma)-\Pi_e\left(a'_{s,h},c,\gamma\right)+K_h\right]$.

To obtain the formula for $\hat{c}_{1,h}\equiv c_{1,h}^U/c_{b,h}^D$ in eq. (21) we must derive

$$c_{1,h}^U=\left[\frac{1+u_h(\gamma)g_h}{1+u_h(\gamma)}\right]^{\frac{1}{\sigma-1}}\left[\frac{a_{1,h}}{(1-\beta_h)K_h}\right]^{\frac{1}{\sigma-1}}\tag{49}$$

and then combine it with the definitions for $a_{1,h}$, $c_{b,h}^D$, g_h in eq.(23) and U_h in eq. (22).

We derive eq. (49) similarly to (46) except the worst case for domestic is the low protection state so instead of eq. (47) we use

$$\mathbb{E}_{s=0,T}V_s^{h\prime}=\lambda_{00}\left[\beta_h\mathbb{E}_{s=0,T+1}V_s^{h\prime}-\pi(a_{s=0,T+1,h},c)+K_h(1-\beta_h)\right]\quad\text{if } c\le c_{s,h}^U\tag{50}$$

Solving forward we obtain $\mathbb{E}_{s=0,0}V_s^{h\prime}=-\lambda_{00}\sum_{t=0}^{\infty}(\beta_h\lambda_{00})^t\pi(a_{s=0,t+1},h,c)+\frac{\lambda_{00}}{1-\beta_h\lambda_{00}}K_h(1-\beta_h)$. Replacing in $\mathbb{E}_1V_s^{h\prime}$ using $\lambda_{00}=1$ and simplifying we obtain

$$\begin{aligned}\mathbb{E}_1V_s^{h\prime}&=\lambda_{10}\left[\beta_h\mathbb{E}_{0,0}V_s^{h\prime}-\pi(a_{0,0h},c)+K_h(1-\beta_h)\right]\\ \mathbb{E}_1V_s^{h\prime}&=\frac{\lambda_{10}}{1-\beta_h}\left[K_h(1-\beta_h)-(1-\beta_h)\left(\sum_{t=0}^{\infty}\beta_h^t\pi(a_{0,t,h},c)\right)\right]\end{aligned}\tag{51}$$

The marginal domestic firm in $s=1$ satisfies $V_1^h\left(c_{1,h}^U\right)=0$, which we use to solve for $c_{1,h}^U$ similarly to the export cutoff but using eq. (51) instead of (48)

$$\begin{aligned}\beta_h\mathbb{E}_1V_1'\left(c_{1,h}^U\right)-\pi\left(a_{1,h},c_{1,h}^U\right)+K_h(1-\beta_h)&=0\\ a_{1,h}\left(c_{1,h}^U\right)^{1-\sigma}\left[1+u_h(\gamma)(1-\beta_h)\sum_{t=0}^{\infty}\beta_h^t\left(\frac{a_{0,t,h}}{a_{1,h}}\right)\right]&=K_h(1-\beta_h)(1+u_h(\gamma))\end{aligned}$$

where $u_h\equiv\frac{\beta_h\lambda_{10}}{1-\beta_h\lambda_{00}}$ and we obtain eq. (49) by using $\frac{a_{0,t,h}}{a_{1,h}}=\left(\frac{P_{0,t}}{P_1}\right)^{\sigma-1}$, g_h from eq.(23) and solving for $c_{1,h}^U$.

B Data and Estimation Appendix

B.1 Data sources and definitions

• *Change in ad valorem Tariffs* $\Delta \ln \tau_V$*:* Log change in 1 plus the statutory ad valorem MFN tariff rate aggregated to the HS6 level between 2005 and 2000. Source: TRAINS via WITS.
• *Change in AVE Tariffs* $\Delta \ln \tau_V$*:* Log change in 1 plus the ad valorem 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* τ_{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 $1 - \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-\sigma}$ computed using year 2000 column 2 and MFN tariff rates.
• *Change in Transport Costs* $\Delta \ln D_V$*:* 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)
• *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

Our policy data for the U.S. and other third countries used in some of the robustness checks have to be concorded over time to the trade data.To do so, we use the published U.N. Statistics Division concordances to map the HS 2002 into the HS 1996. This provides uniformity across all tariff and trade data sources at the 6 digit level. At more disaggregated levels with the NBER trade flow data, we use the method described by Pierce and Schott (2012) to match and combine the 10-digit level import flows over time. We then aggregate up to the 6 digit level of HS 1996 when constructing price indices or product variety counts as needed.[72]

B.2 Expenditure share, import penetration and risk counterfactuals

Import penetration in manufacturing is Chinese imports over U.S. expenditure on manufacturing, $R_{Ch,t}/E_t$. We define total manufacturing expenditure, $E = \mu L$ in the model, as total manufacturing shipments less net manufacturing exports, E_t =Manuf. Shipments$_t$ $- Exports_t + Imports_t$. We compute $\mu = 0.86$ as the share of manufacturing in total expenditure on tradables (=Gross Output - 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 in year t, we follow the exact same steps as for the baseline year (2005). Thus we employ the observed import penetration for each year $t = 2002 \ldots 2010$, adjust it to account for the change in tariffs relative to 2000, and compute the change in imports due to TPU. We use this to compute the counterfactual imports from China normalized by expenditure, $R^{CF}_{Ch,t}/E_t$, 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 for each industry. We adjust import penetration to the level implied by the resulting weighted mean tariff of $\overline{\tau} = 1.14$. The procedure uses the

[72] Some disaggregated trade flows are reassigned across multiple 6 digit HS codes over time. Because these codes can't be tracked longitudinally without further and sometimes arbitrary aggregation of certain 6 digit HS codes, we drop them for all years from 1996-2006.

2005 import penetration to compute the price elasticity to a tariff change. With the model quantities all adjusted to their levels at the mean of the tariff distribution, we can then compute the GE effect on exports, entry, and other quantities around the mean. We follow the same procedure to compute the GE solution over a grid of counterfactual initial applied tariff regimes in Figure 8.

B.3 Sunk cost estimation

Approach

In the model, uncertainty only has an effect for industries with positive sunk costs. To empirically identify those industries we explore variation in export persistence across countries exporting to the U.S. A standard approach (cf. Roberts and Tybout, 1997) is to use firm-level data to estimate a probability model where, after conditioning on firm characteristics to capture their current incentive to participate, any correlation with lagged participation provides evidence of sunk costs. Our objective is not to estimate the magnitude of sunk costs in each industry but simply to determine which subset is more likely to have sunk costs and then use it to test if uncertainty has stronger effects in those industries.

More formally, let the export participation variable be $Y_{vct} = \{0, 1\}$ for firm v from export country c at t. We define an indicator for a sunk versus fixed export cost industry: $\kappa_V = 1$ if $K_V > 0$ and $f_V = 0$ and $\kappa_V = 0$ if $K_V = 0$ and $f_V > 0$. Clearly there are country and time dimensions to these costs, which we are ignoring in the exposition. Denote the equilibrium industry threshold for new exporters from country c at t, i.e. those with $Y_{vc,t-1} = 0$, as $c_{ct}(\kappa_V)$. This is the cutoff we solved for in the model when $\kappa_V = 1$; for an industry with fixed costs we would obtain $c_{ct}(\kappa_V = 0) = \left[\frac{a_{Vct}}{f_{Vc}}\right]^{\frac{1}{\sigma-1}}$. The participation equation for a firm with cost parameter c_{vct} in period t under fixed export costs is independent of prior participation and given by

$$Y_{vct}(\kappa_V = 0) = \begin{cases} 1 \text{ if } c_{vct} \leq c_{ct}(\kappa_V) \\ 0 \ \text{ otherwise.} \end{cases} \tag{52}$$

Alternatively, under sunk costs, a firm will export in the current period if (i) its marginal cost parameter satisfies the current cutoff condition $c_{vct} \leq c_{ct}(\kappa_V)$, or; (ii) its marginal cost exceeds the cutoff but it exported in the previous period $(c_{vct} > c_{ct}(\kappa_V) \wedge Y_{vc,t-1} = 1)$. The participation equation is

$$Y_{vct}(\kappa_V = 1) = \begin{cases} 1 \text{ if } c_{vct} \leq c_{ct}(\kappa_V) \vee Y_{vc,t-1} = 1 \\ 0 \text{ otherwise.} \end{cases} \tag{53}$$

We capture firm participation by using HS-10 product data over 1996-2000 for a set of exporters to the U.S. market. Each industry V is composed of a group of HS-10 categories, denoted by $\tilde{V} \in V$. Within each country$\times$HS-10 category there is a subset of firms and we denote the cost of the most productive one by $c_{\tilde{V}ct}$. We note three points about mapping from the model to the product data. First, even if the productivity distribution at the HS-6 level is unbounded, it is possible to have certain HS-10 products where $c_{\tilde{V}ct} > c_{ct}(\kappa_V)$ so no trade would be observed under fixed costs (or under sunk costs if $Y_{vc,t-1} = 0$). Thus the variation in export participation that we explore at the HS-10 level is consistent with the TPU augmented gravity equation we derived. Second, the model does not assume any correlation between the product category $\tilde{V} \in V$ that a given firm v produces and that firms' productivity. Thus we treat each set of firms $v \in \tilde{V}$ as a random partition of the productivity distribution of its respective HS-6 industry and model the minimum cost as an unobserved parameter: $c_{\tilde{V}ct} = c_{\tilde{V}} c_{ct} \exp(\varepsilon_{\tilde{V}ct})$ where $\varepsilon_{\tilde{V}ct}$ is a random error term.

Defining the latent variable $z_{\tilde{V}ct}(\kappa_V) \equiv \ln(c_{ct}(\kappa_V)/c_{\tilde{V}ct})$ we can write the HS-10 counterpart of (52) as $T_{\tilde{V}ct}(\kappa_V = 0) = 1$, if $z_{\tilde{V}ct}(\kappa_V = 0) \geq 0$ and 0 otherwise; and for (53) we have $T_{\tilde{V}ct}(\kappa_V = 1) = 1$, if $z_{\tilde{V}ct}(\kappa_V = 1) \geq 0 \vee T_{\tilde{V}c,t-1} = 1$ and 0 otherwise.

Identification and estimation

The theoretical model and the assumption made about $c_{\tilde{V}ct}$ allows us to write the latent variable as a function of fixed effects and an error term, $z_{\tilde{V}ct}(\kappa_V) = \alpha_{Vct} + \alpha_{\tilde{V}} + \varepsilon_{\tilde{V}ct}$, which applies whether $\kappa_V = 0, 1$.

The country-year-industry effects capture all the factors the theory allows for in the economic conditions variable, a_{Vct}, that enters $c_{ct}(\kappa_V)$, e.g. it subsumes the aggregate U.S. expenditure and price index effects, allows for (HS-6) industry tariffs, transport and other export costs to differ across countries. If a country is particularly productive in a given industry then this is controlled for by α_{Vct}. We allow for the possibility that certain products contain more (or less productive) firms via the HS-10 effect, $\alpha_{\tilde{V}}$.

We estimate a linear probability model to handle the large set of fixed effects:

$$T_{\tilde{V}ct} = b_V^{sunk} T_{\tilde{V}ct-1} + b_{V,96} T_{\tilde{V}c,96} + \alpha_{Vct} + \alpha_{\tilde{V}} + \epsilon_{\tilde{V}ct} \text{ for each } V.$$

To address any remaining unobserved heterogeneity in initial conditions at the HS10-country level we also control for the export status in the first year of the sample, $T_{\tilde{V}c,96}$. In order to identify b_V^{sunk} there must exist sufficient changes in trade status in an industry and some firms that are exporting even though their marginal cost is above the current cutoff. This requires us to have a sufficiently large number of time-country-HS10 observations. We restrict the countries to exclude China and the time period to the one prior to China's WTO accession, 1996-2000, to avoid these results being affected by China's export boom.[73] Thus to increase the number of observations and better identify b_V^{sunk} we estimate the model at the HS-4 level. Doing so implicitly restricts the HS-6 industries in each HS-4 to have similar parameters. This restriction is more likely to be met by a group of countries that face similar trade protection, so we estimate the model using U.S. imports from non-preferential partners other than China.

Estimates

The coefficient of interest is b_V^{sunk}. The null hypothesis in a model with fixed costs and no sunk costs is that $b_V^{sunk} = 0$; we interpret $b_V^{sunk} > 0$ as evidence for the presence of sunk costs. Figure A3 plots the t-statistics against the estimated coefficients. The results appear reasonable along a couple of dimensions. First, only 29 of 1,084 estimates are negative and all but two of those negative estimates are insignificantly different from 0. Second, the increase in the probability of exporting due to lagged exporting is always lower than one, the maximum is 0.81.

Figure A3 also shows there is heterogeneity in persistence across industries. This is useful in providing us with a ranking that allows us to distinguish between industries according to how likely they are to have sunk costs. To do so we rank industries by the persistence coefficients' t-statistic; those industries where we reject fixed costs (no persistence) with a higher confidence level are those we classify as having relatively higher sunk costs.[74] About three quarters of the industries have a t-statistic above 2.58 (around 1% significance level) and two thirds are above 3.09 (around 0.2% significance), represented by the red line.

We match these estimates to the HS-6 sample used in table 6 and define $\tilde{\kappa}_V = 1$ for those industries with t-statistics in the top two terciles of that sample as more likely to have sunk costs than those in the bottom tercile, $\tilde{\kappa}_V = 0$. There is no obvious metric to compare our estimates to since there is no accepted measure of export sunk costs for this large a set of industries. However, we can ask if the estimates are informative about persistence and thus sunk costs for China. To do so we note that one of the underlying assumptions of the estimation is that sunk export costs have an important industry dimension, which is similar across exporters to the same destination. If this is true then we expect to find a significantly higher autocorrelation in export status for the subset of industries that we identify as higher sunk cost for countries *not* used in the estimation. The more relevant for us is China's exports to the U.S., for which we obtain:

$$T_{\tilde{V}china,t} = \underset{(.009)}{.63}\, T_{\tilde{V}china,t-1} + \underset{(.085)}{.29} \text{ for all } \tilde{\kappa}_V = 1$$

$$T_{\tilde{V}china,t} = \underset{(.023)}{.55}\, T_{\tilde{V}china,t-1} + \underset{(.022)}{.41} \text{ for all } \tilde{\kappa}_V = 0$$

Thus, China's lagged exporting in a product has a significant effect on current exporting and, more importantly, that effect is stronger for industries that our procedure identifies as high sunk cost. We obtain

[73] Because we include a lagged term in the dependent variable, the year 1996 is dropped as an outcome year.

[74] The number of observations is not the same across V but they are large enough in each of them such that higher t-statistics translate into higher confidence intervals.

a similarly significant difference in persistence if we re-run these specifications while using HS-4 effects to control for the possibility that China may be more productive in those industries where $\tilde{\kappa}_V = 1$ (coefficient is 0.56) than $\tilde{\kappa}_V = 0$ (0.49). These results hold whether we focus on $t = 2000$, as described, or we include additional years.

B.4 Industry Price Indices: Measurement, Predictions, and Aggregation

We describe the measurement and model predictions for the following change in ideal prices in an industry across two periods t and $t-5$:

$$\Delta \ln P_{V,x} \equiv \ln \left[\frac{\int_{\Omega^x_{tV}} (p_{tv})^{1-\sigma}}{\int_{\Omega^x_{t-5V}} (p_{t-5v})^{1-\sigma}} \right]^{1/1-\sigma} \tag{54}$$

Measurement

Feenstra (1994) shows that exact changes in the CES ideal price index can be computed as a function of weighted changes in the prices of continuing varieties, and a term accounting for changes in varieties. Applying the derivation to eq. (54) we obtain

$$\Delta \ln P_{V,x} = \sum_{v \in \Omega^{cont}_{V,x}} w_{v,t} \ln\left(\frac{p_{v,t}}{p_{v,t-5}}\right) + \ln\left(\frac{\psi_{V,t}}{\psi_{V,t-5}}\right)^{1/(\sigma-1)} \tag{55}$$

where $\Omega^{cont}_{V,x}$ is the set of imported varieties in industry V traded in both periods, $p_{v,t}$ is their consumer price in t and $w_{v,t}$ are ideal variety share weights defined by

$$w_{v,t} \equiv \frac{(s_{v,t} - s_{v,t-5}) / (\ln(s_{v,t}) - \ln(s_{v,t-5}))}{\sum_{v \in \Omega^{cont}_{V,x}} ((s_{v,t} - s_{v,t-5}) / (\ln(s_{v,t}) - \ln(s_{v,t-5}))}$$

$$s_{v,t} \equiv \frac{p_{v,t} q_{v,t}}{\sum_{v \in \Omega^{cont}_{V,x}} p_{v,t} q_{v,t}} \; ; \; s_{v,t-5} \equiv \frac{p_{v,t-5} q_{v,t-5}}{\sum_{v \in \Omega^{cont}_{V,x}} p_{v,t-5} q_{v,t-5}}$$

The variety adjustment measures the change in the expenditure share of continuing varieties.

$$\psi_{V,t} \equiv \frac{\sum_{v \in \Omega^{cont}_{V,x}} p_{v,t} q_{v,t}}{\sum_{v \in \Omega_{V,x,t}} p_{v,t} q_{v,t}} \; ; \; \psi_{V,t-5} \equiv \frac{\sum_{v \in \Omega^{cont}_{V,x}} p_{v,t-5} q_{v,t-5}}{\sum_{v \in \Omega_{V,x,t-5}} p_{v,t-5} q_{v,t-5}}$$

We follow Broda and Weinstein (2006) in defining a variety as an HS-10 product by country observation. Our calculation differs from theirs in three ways. First, we assume σ is similar across industries. Second, we compute the change for $t = 2005$. Third, we compute separate sub-price indices for China (and other U.S. trading partners), which can be aggregated across industries (as done in Broda and Weinstein, eq. 12) and similarly across countries. More specifically, we do the following:

1. Concord HS-10 data over time using an algorithm similar to Pierce and Schott (2010) modified to account for details of the tariff classification.

2. Compute unit values at HS-10 for each year if quantity is available and $\Delta \ln p_v$ if v is traded in both periods and its quantity is reported in the same units.

3. Define $\mathbf{V}^{cont}$ as the set of industries with at least one measured variety price change, $\Delta \ln p_{v \in V} \neq \emptyset$, and the associated set of continuing varieties, $\Omega^{cont}_{V,x}$ for each $V \in \mathbf{V}^{cont}$. The baseline defines V at the HS6 level.

4. Compute $\psi_{V,t}, \psi_{V,t-5}$, $w_{v,t}$ and use eq.(55) to obtain $\Delta \ln P_{V,x}$ for each $V \in \mathbf{V}^{cont}$.

Sample selection and measurement error:

Using the procedure above the number of HS-6 industries where $V \in \mathbf{V}^{cont}$ and for which the variables in the gravity estimation are available is $n = 2714$. Thus we can compute ideal price changes for 85% of the HS-6 export sample (2714/3211) either because the index is not defined or because of unavailability of quantity data. Thus in some of the robustness tests we define V at the HS-4 level, which ensures that a smaller fraction of industries is dropped since $\Omega^{cont}_{HS6,x} \subseteq \Omega^{cont}_{HS4,x}$.

We measure price changes with error by using changes in average unit values. Given this is our dependent variable we treat this error as random across industries. If unit values are poorly measured in some sectors then the specification with sector effects control for it. Nonetheless, there are outliers both at the top and bottom (about 6.5% of the sample is mild outliers and 3% severe, i.e. +/- 3 times the interquartile range). To minimize their potential effect we trim the top and bottom 2.5 percentiles leaving 2579 observations.

Predictions

To obtain the estimating equation (12) we use the price change defined in eq. (54) and the derivation in eq.(42). Allowing for exogenous changes in export costs other than tariffs in eq.(42) the log change in the import index in a temporary state s relative to a deterministic baseline b is

$$\ln\left(\frac{P_{sV,x}}{P_{bV,x}}\right) = \ln\left(\frac{\tau_{sV}}{\tau_{bV}}\frac{d_{sV}}{d_{bV}}\right) + \left(1 - \frac{k}{\sigma-1}\right)\ln\left(\frac{c^U_{sV}}{c^D_{bV}}\right)$$

The estimation uses $\Delta \ln x = \ln \frac{x_{bV}}{x_{sV}}$ since the post period is the deterministic baseline, and $s = 1$. Using this and the generalized version of the formula in eq. (18): $c^U_{1V}/c^D_{0V} = U(\omega g, \gamma) \times \left(\frac{a_{1V}}{a_{0V}}\right)^{\frac{1}{\sigma-1}}$, we obtain:

$$\begin{aligned}
\Delta \ln P_{V,x} &= \Delta \ln(\tau_V) + \Delta \ln(d_V) + \left(1 - \frac{k}{\sigma-1}\right)\left[\frac{1}{\sigma-1}\Delta \ln(a_V) - \ln U_V\right] \\
&= \left(1 - \frac{k}{\sigma-1}\right)(-\ln U_V) + \left(\frac{\sigma k}{\sigma-1} - 1\right)\frac{1}{\sigma-1}\Delta \ln \tau_V + \frac{k}{\sigma-1}\Delta \ln d_V \\
&\quad + \left(1 - \frac{k}{\sigma-1}\right)\Delta \ln\left(PE^{\frac{1}{\sigma-1}}\right)
\end{aligned} \tag{56}$$

where the second equality uses $a_V \equiv (\tau_V \sigma)^{-\sigma}\left((\sigma-1)P/d_V\right)^{\sigma-1} E$. The last term captures any aggregate changes, which are endogenous to the policy change in the general case, or exogenous in the small exporter case. The empirical counterpart in eq. (12) reflects an error term due to potential measurement problems in the price indices, as described above, and possibly from measuring d_V with D_V, i.e. with freight and insurance information alone.

Aggregation

When aggregating industry import price index changes using the $P_{V,x}$ constructed from the data we use

$$\Delta \ln P_x \equiv \sum_V w_{Vt,x}\Delta \ln P_{V,x}, \text{ where } w_{Vt,x} \equiv \frac{(s_{V,t} - s_{V,t-5})/\left(\ln(s_{V,t}) - \ln(s_{V,t-5})\right)}{\sum_V \left((s_{V,t} - s_{V,t-5})/(\ln(s_{V,t}) - \ln(s_{V,t-5})\right)}.$$

B.5 Entry: Measurement and Predictions

Predictions

The model predicts the growth in imported varieties, $\Delta \ln n_V$, after switching from a temporary policy state, 1, to a permanent one, 0, is

$$\Delta \ln n_V = k \ln c^D_{0V}/c^U_{1V} = -k \ln U_V + \frac{k}{\sigma-1}\Delta \ln (a_V)$$
$$= k(-\ln U_V) - \frac{\sigma k}{\sigma-1}\Delta \ln \tau_V - k\Delta \ln d_V + \frac{k}{\sigma-1}\Delta \ln \left(PE^{\frac{1}{\sigma-1}}\right) \tag{57}$$

where the second quality in the first line uses the generalized version of the formula in eq. (18): $c^U_{1V}/c^D_{0V} = U(\omega g, \gamma) \times \left(\frac{a_{1V}}{a_{0V}}\right)^{\frac{1}{\sigma-1}}$. The second line uses $a_V \equiv (\tau_V \sigma)^{-\sigma}((\sigma-1)P/d_V)^{\sigma-1} E$ and allows for any aggregate changes, which are endogenous to the policy change in the general case, or exogenous in the small exporter case. The empirical counterpart in eq. (13) reflects an error term due to potential measurement problems in the the number of varieties, as described below, and measuring d_V with D_V, i.e. with freight and insurance information alone.

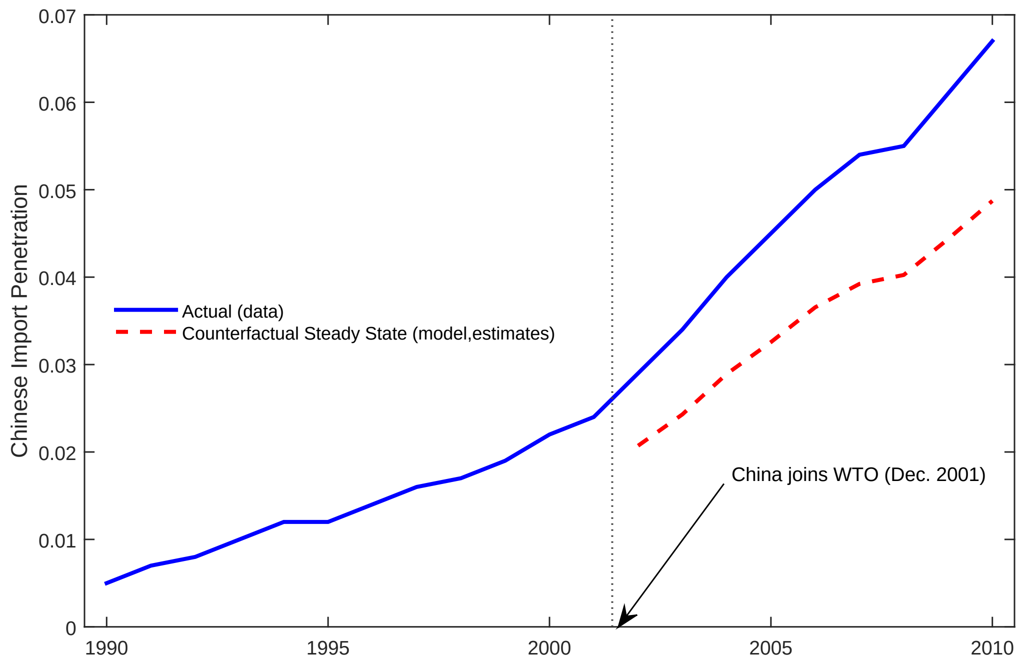
Measurement and estimation

We measure varieties as HS-10 products by country and thus variety growth is the growth in traded HS-10 within an industry V. The growth is censored for any HS-6 industries where all HS-10 categories are traded in both periods and it provides no information about variety growth. Thus using the full sample to estimate eq. (13) yields attenuated estimates of the coefficients and we can minimize it by focusing on the uncensored sample, as shown in Table 8.

Moreover, under certain conditions we can identify the coefficients implied by eq. (57) up to a factor, $\nu' \in [0,1]$. Assume there is a continuous, increasing, differentiable function $\nu(\cdot)$ that maps varieties to product counts: $\ln(pcount_{sV}) = \nu(\ln n_{sV})$. If there was only one firm in an HS-6 industry and it produced a single variety then we would observe one traded HS-10 within that 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 $\nu(\ln n_{sV} = 0) = 0$ and an upper bound $\ln(pcount^{\max}_{tV}) = \nu(\ln n_{hV})$ for all $\ln n_{tV}$ at least as high as $\ln n_{hV}$—the (unobserved) threshold where all HS-10 product categories in an HS-6 industry have positive values. If we assume product counts and true varieties are continuous, then $\nu' \geq 0$ for $n_V \in (0, n_{hV})$ 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. If we log linearize the equation of product counts around $\ln n_{t-5V}$ the change in products between t and and $t-5$ is $\Delta \ln(pcount_V) \approx \nu'(\ln n_{s-1V})\Delta \ln n_V$. Therefore, if we use $\Delta \ln(pcount_V)$ as a proxy for $\Delta \ln n_V$ we can identify the coefficients in eq. (57) up to a factor, $\nu'(\ln n_{s-1V})$, if that factor is similar across industries.

Figures and Tables

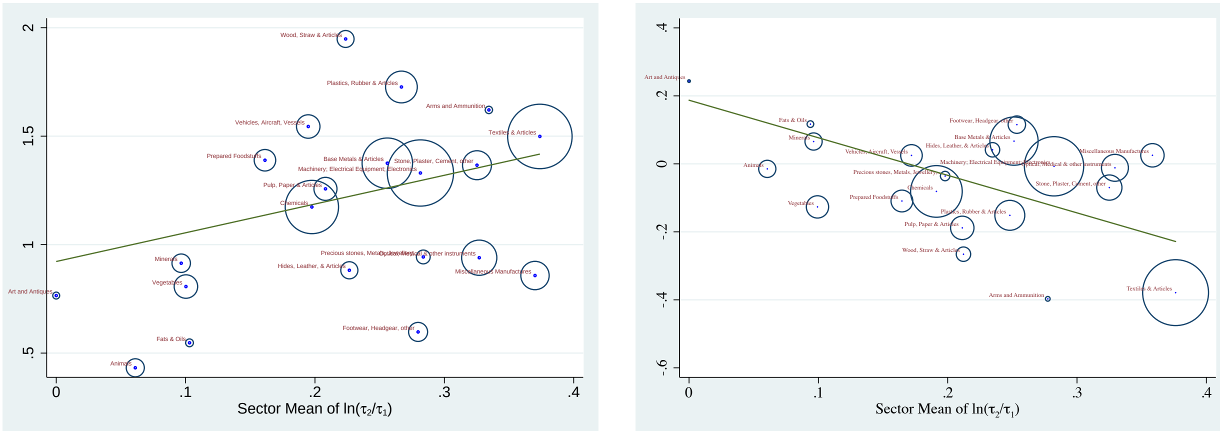
Figure 1: Chinese Import Penetration in U.S. – Actual vs. Counterfactual under Policy Uncertainty.



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

Figure 2: Sector level Chinese export and price index growth vs. initial uncertainty

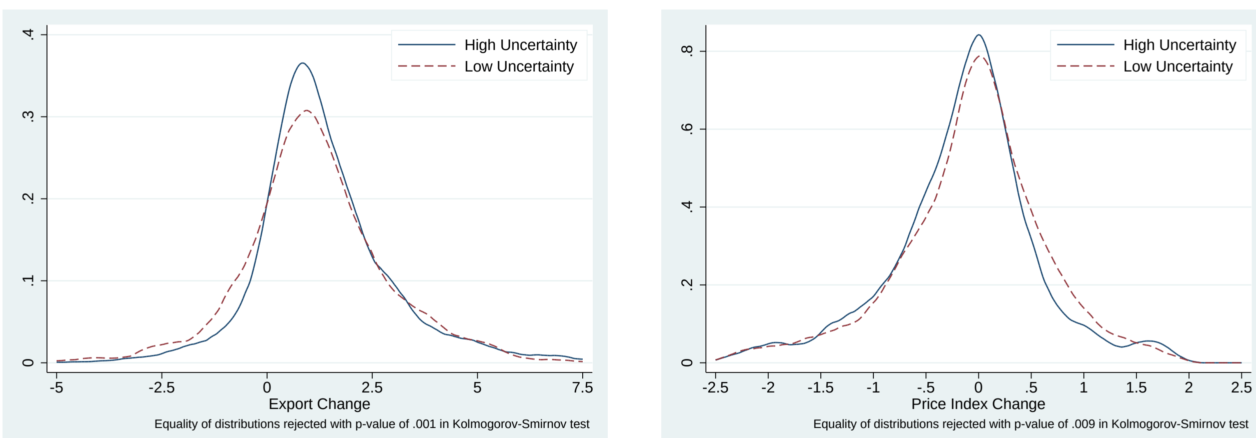
(a) Exports ($\Delta \ln$) **(b) Price Index ($\Delta \ln$)**



Notes: Simple means within sector of export and price index change vs means of initial uncertainty measured by $\ln(\tau_{2V}/\tau_{1V})$ where τ_{2V} and τ_{1V} are the column 2 and MFN tariff factors in 2000. Circles are proportional to the number of observations used as weights in the linear fit represented.

Figure 3: Chinese export and price index growth in High vs Low Uncertainty Industries

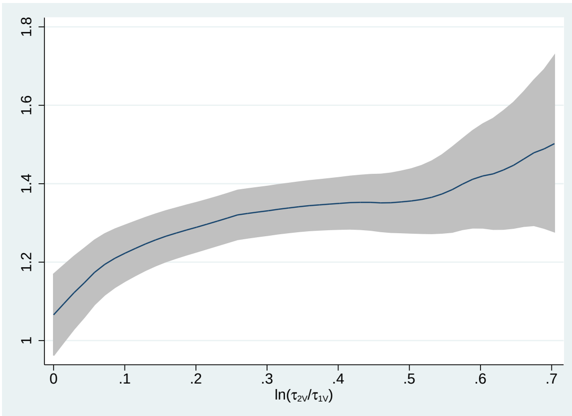
(a) Exports ($\Delta \ln$) **(b) Price Index ($\Delta \ln$)**



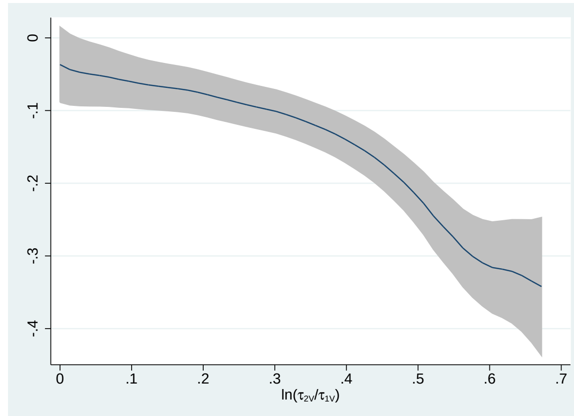
Notes: Epanechnikov kernel density estimates. High uncertainty defined as the top two terciles of ratio U.S. Column 2 to MFN tariff. Low uncertainty is the bottom tercile. See text for details of price index calculation.

Figure 4: Chinese export and price index growth ($\Delta \ln$) vs. initial policy uncertainty

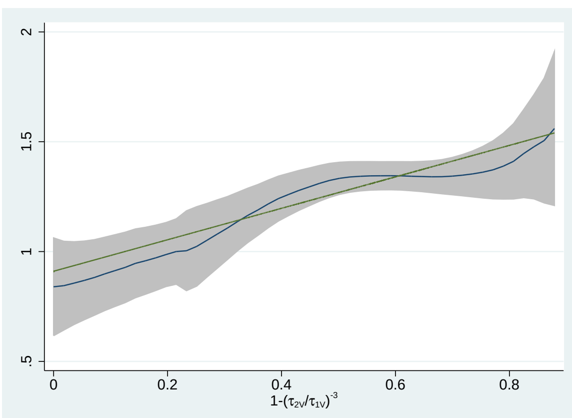
(a) Exports–local polynomial fit



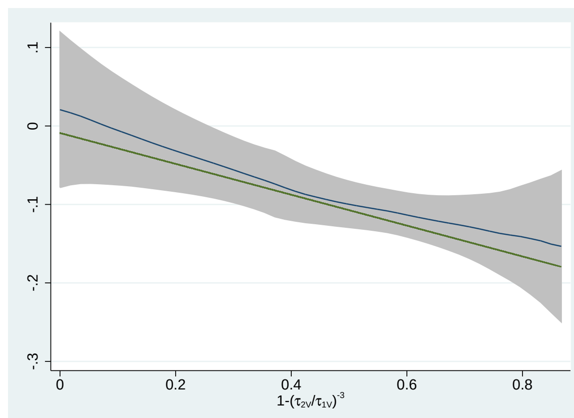
(b) Price index–local polynomial fit



(c) Exports–semi-parametric fit

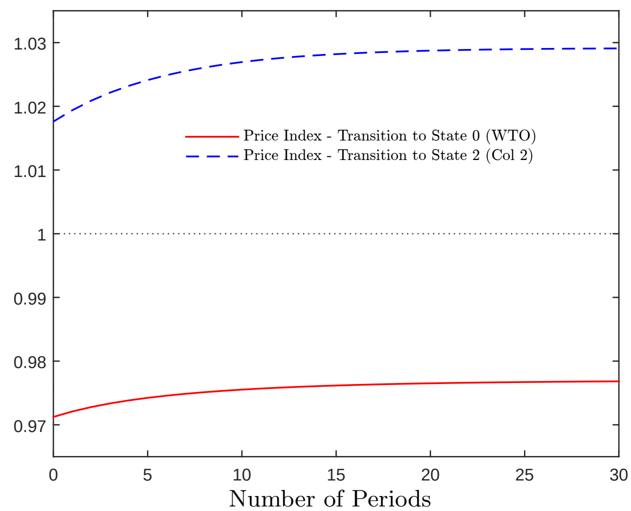


(d) Price index–semi-parametric fit



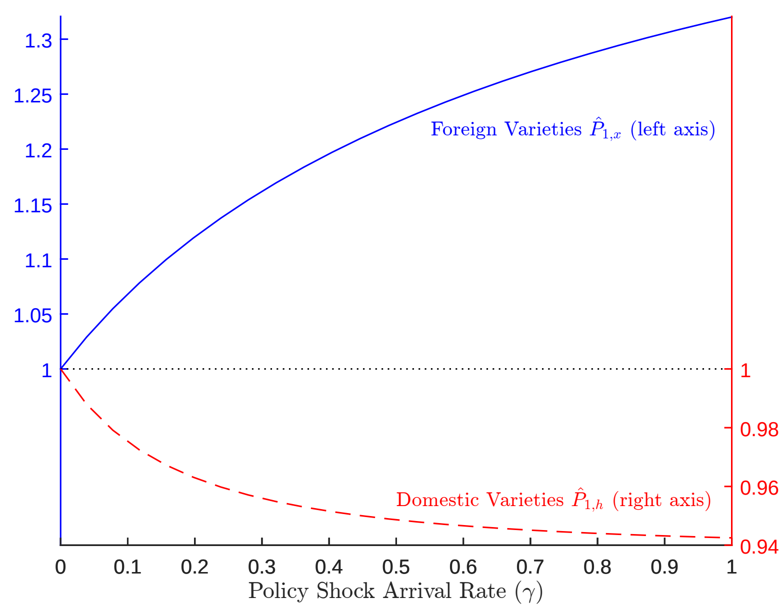
Notes: Panels (a) and (b) are a local polynomial fit on $\ln(\tau_{2V}/\tau_{1V})$ where τ_{2V} and τ_{1V} are the column 2 and MFN tariff factors in 2000. Panels (c) and (d) regress log export and log price index growth on changes in transport costs, tariffs, and on sector dummies. The linear fit uses OLS and also includes $-(\tau_{2V}/\tau_{1V})^{-3}$, 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 $1-(\tau_{2V}/\tau_{1V})^{-3}$ for ease of comparison with the uncertainty variable used in the baseline OLS regressions.

Figure 5: Price index transition dynamics from intermediate to high or low protection state



Notes: General equilibrium solution of the model for estimated and assumed parameters in Table A9 and $\gamma = 0.248$

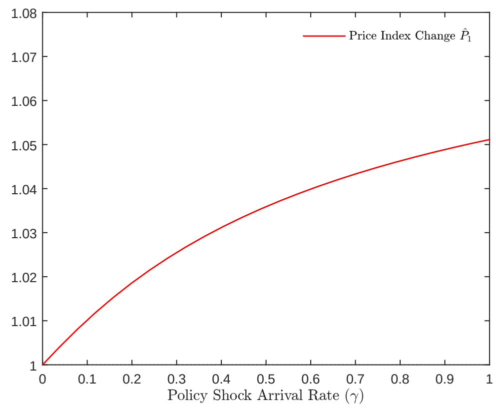
Figure 6: Domestic and Foreign Variety Price Indices (Counterfactual introduction of U.S. TPU on all partners)



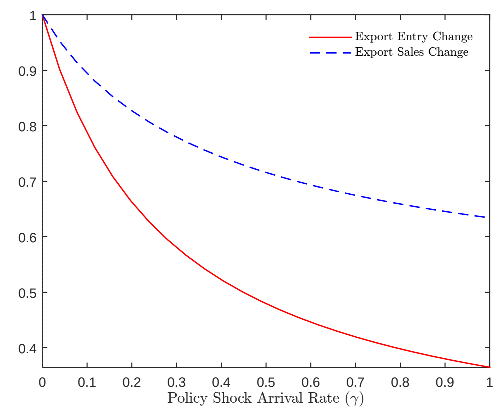
Notes: General equilibrium solution of the model for each variable's growth relative to deterministic baseline if the U.S. introduces TPU in 2005. Solutions computed from estimated and assumed parameters in Table A9.

Figure 7: Aggregate Price Index, Sales and Entry vs. γ (Counterfactual introduction of U.S. TPU on all partners)

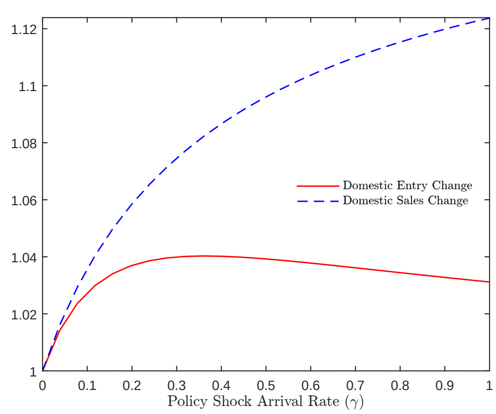
(a) Aggregate Price Index



(b) Export Sales and Entry



(c) Domestic Sales and Entry

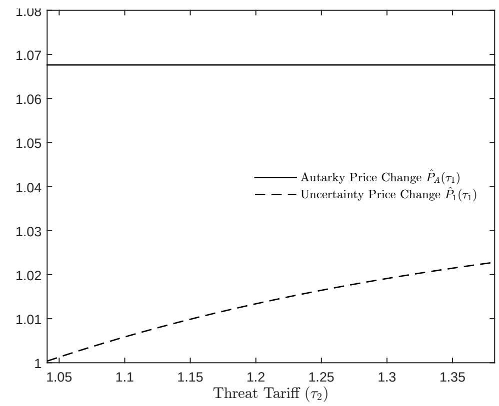


Notes: General equilibrium solution of the model for each variable's growth relative to deterministic baseline if the U.S. introduces TPU in 2005. Solutions computed from estimated and assumed parameters in Table A9.

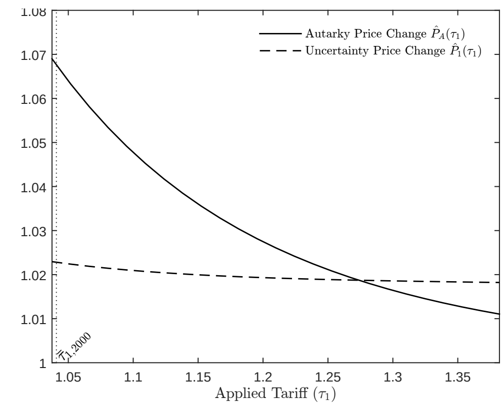
Figure 8: Aggregate Price, Sales and Entry vs. Alternative Applied and Threat Tariffs (Counterfactual introduction of U.S. TPU on all partners)

Threat Tariff (τ_2) | **Applied Tariff (τ_1)**

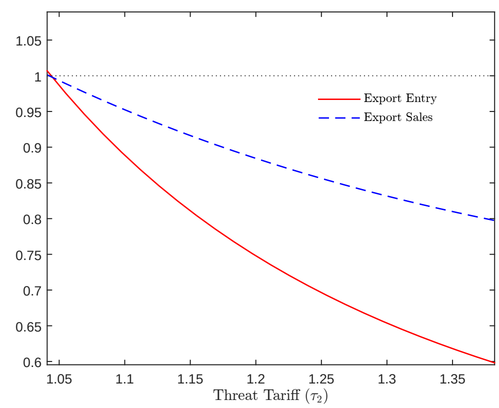
(a) Autarky and Uncertainty Price Index



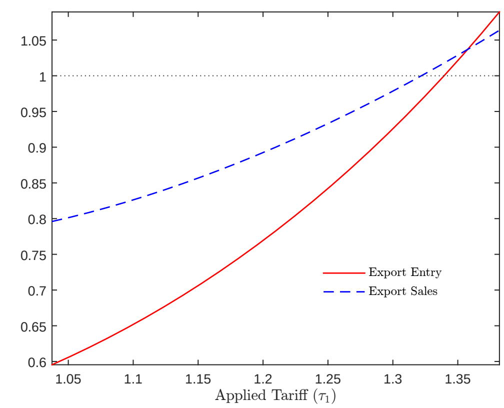
(b) Autarky and Uncertainty Price Index



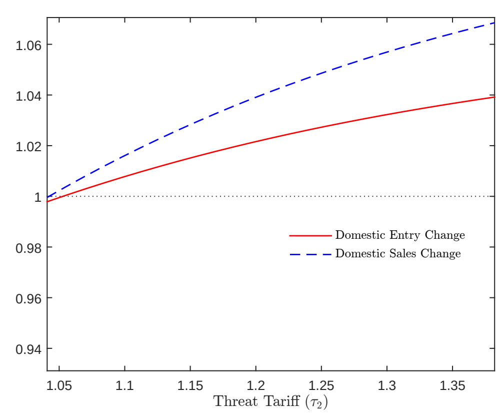
(c) Export Sales and Entry



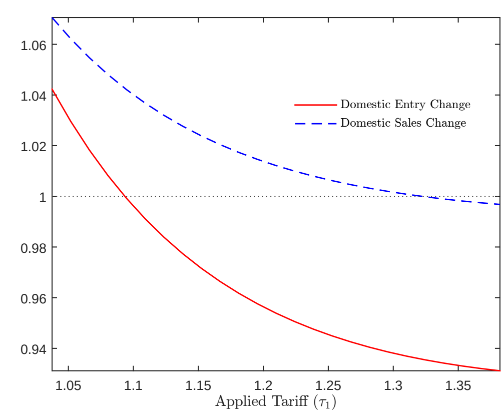
(d) Export Sales and Entry



(e) Domestic Sales and Entry



(f) Domestic Sales and Entry



Notes: General equilibrium solution of the model for each variable's growth relative to deterministic baseline if the U.S. introduces TPU in 2005. Solutions computed from estimated and assumed parameters in Table A9. For figures (a),(c), and (e), the threat tariff (τ_2) on the x-axis ranges from the simple mean of the observed MFN tariff in 2000 of 1.041 to the Column 2 tariffs in 2000 of 1.38. For figures (b), (d), and (f), the applied tariff (τ_1) on the x-axis ranges from the simple mean of the observed MFN tariff in 2005 of 1.038 to the Column 2 tariffs in 2000 of 1.38. Aggregate import weights and import penetration are adjusted for the counterfactual τ_1 relative to observed values in 2005. For (a) the price change from increasing uncertainty or autarky is computed relative to a fixed applied $\tau_1 = 1.041$. For (b) the autarky price change is computed over $\tau_1 \in [1.038, 1.38.]$

Table 1: Summary statistics by pre-WTO policy uncertainty

	Uncertainty		
	Low	**High**	**Total**
Chinese export value growth to U.S. (Δln, 2005-2000)	1.16	1.35***	1.28
	[1.772]	[1.617]	[1.675]
Chinese export price index growth (Δln, 2005-2000)[a]	-0.07	-0.14**	-0.11
	[0.700]	[0.690]	[0.694]
Chinese export variety growth (Δln, 2005-2000)[b]	0.27	0.35***	0.32
	[0.457]	[0.409]	[0.432]
MFN tariff (ln), 2000	0.027	0.041	0.036
	[0.036]	[0.046]	[0.043]
Column 2 tariff (ln), 2000	0.159	0.393	0.311
	[0.096]	[0.116]	[0.156]
Ratio of Col 2 to MFN tariff	1.145	1.429	1.330
	[0.090]	[0.140]	[0.184]
Potential profit loss if MFN revoked (pre WTO)	0.303	0.637	0.520
	[0.176]	[0.086]	[0.203]
MFN tariff (Δln)	-0.002	-0.004	-0.003
	[0.007]	[0.010]	[0.009]
Transport costs (Δln)	-0.010	-0.002	-0.005
	[0.100]	[0.079]	[0.087]
Observations	1124	2087	3211

Notes:
Simple means with standard deviations in brackets. Low: subsample of industries in the bottom tercile of pre-WTO uncertainty (ranked by τ_2/τ_1); High refers to the rest of the sample. Total includes the full sample used in baseline Table 2.
*** 1% significance level and ** 5% significance level for difference of growth between low and high subsamples.

a. Total observations for price index change are 2579. High and low bins are defined on the baseline sample.
b. Total observations for variety growth (number of traded HS-10 varieties with an HS-6 industry) are 1051 and exclude industries that are censored above, i.e all varieties are traded in 2000 and 2005. High and low bins are defined on the baseline sample.

Table 2: Chinese Export Growth (2000-2005, U.S., Δln) — Baseline Estimates

	1	2	3	4
Uncertainty Pre-WTO	0.743***	0.791***	0.716***	0.734***
[+]	[0.154]	[0.150]	[0.186]	[0.184]
Change in Tariff (Δln)	-9.967**	-4.340***	-7.356	-4.250***
[-]	[4.478]	[0.676]	[5.060]	[0.677]
Change in Transport Costs (Δln)	-2.806***	-2.893***	-2.795***	-2.833***
[-]	[0.455]	[0.450]	[0.456]	[0.451]
Constant	0.851***	0.843***		
	[0.0853]	[0.0850]		
Observations	3,211	3,211	3,211	3,211
R-squared	0.03	n/a	0.05	n/a
Sector fixed effects	no	no	yes	yes
Restriction p-value (F-test)	0.204	1	0.536	1

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 σ=3. All specifications employ OLS and 2 and 4 impose theoretical constraint on tariffs and transport cost coefficients: $b_\tau=b_d(\sigma/(\sigma-1))$. Sectors defined by the 21 HS sections.

Table 3: Chinese Export Growth (2000-2005, U.S., Δln) — Robustness to NTBs

	1	2	3	4	5
Specification:	Baseline	+MFA/TTB	+MFA/TTB +Sector FE	+MFA/TTB +Sector FE IV (NTB)	+MFA/TTB +Sector FE Constrained
Uncertainty Pre-WTO	0.743***	0.679***	0.720***	0.726***	0.744***
[+]	[0.154]	[0.152]	[0.185]	[0.185]	[0.184]
Change in Tariff (Δln)	-9.967**	-8.979**	-8.272	-8.397*	-4.300***
[-]	[4.478]	[4.559]	[5.058]	[5.048]	[0.672]
Change in Transport cost (Δln)	-2.806***	-2.797***	-2.818***	-2.825***	-2.867***
[-]		[0.452]	[0.452]	[0.450]	[0.448]
Change in MFA quota status		-0.188*	-0.313**	-0.313**	-0.304**
		[0.101]	[0.136]	[0.136]	[0.135]
Change in NTB status		-0.944***	-0.974***	-1.309	-0.968***
		[0.317]	[0.330]	[0.904]	[0.330]
Constant	0.851***	0.871***			
	[0.0853]	[0.0845]			
Observations	3,211	3,211	3,211	3,211	3,211
R-squared	0.03	0.04	0.06	0.06	.
Sector fixed effects	no	no	yes	yes	yes
F-stat, 1st Stage	.	.	.	10.21	.
Over-ID restriction (p-value)	.	.	.	0.592	.
Restriction p-value (F-test)	0.204	0.3	0.428	0.414	1

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 imposes theoretical constraint on tariffs and transport cost coefficients: $b_\tau=b_d(\sigma/(\sigma-1))$. 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 σ=3

Table 4: Chinese Export Growth (2000-2005, Δln) — Robustness to unobserved HS-6 export supply shocks

Dependent variable	**Chinese export** growth to:			
	U.S.	**EU-15**	**Japan**	**Pooled**
	1	2	3	4
Uncertainty pre-WTO (U.S.)	0.554***	0.0174	0.208	-
	[0.193]	[0.186]	[0.176]	
Uncertainty pre-WTO (U.S.) x 1(U.S.)				0.428**
				[0.210]
MFN Tariff (Δln)	-6.042	-7.970***	-8.306	-5.080*
	[5.120]	[2.949]	[5.678]	[2.640]
Observations	3,100	3,004	2,723	8,827
R-squared	0.03	0.04	0.08	0.05
Sector Fixed Effects	yes	yes	yes	no
HS6 Fixed Effects	no	no	no	yes
Sector*Country Fixed Effects	no	no	no	yes
Equality of Tariff Coeffs (p-value)				0.122
Equality of EU & Japan Uncertainty Coef. (p-value)				0.261

Notes:
Robust standard errors in brackets for columns 1-3. HS6 product clustered standard errors in column 4. *** p<0.01, ** p<0.05, * p<0.1 Uncertainty pre-WTO is defined as in the baseline US sample. The MFN tariff change is the tariff applied to China by the importing country. Transport cost data for Chinese exports to EU and Japan is unavailable. The pooled sample in column 4 is the subset of HS6 products with trade in 2000 and 2005 for Chinese exports to US matched to export flows to either the EU-15, Japan, or both. Columns 1-3 are the export destination subsets of the pooled sample.

Table 5: U.S. Import Growth (2000-2005, Δln) — Robustness to unobserved HS-6 import demand shocks

	Matched Sample of U.S. import growth from China and all non-Preferential MFN partners		Matched Sample of U.S. import growth from China and Taiwan	
	1	2	3	4
Uncertainty x **1**(China)	0.751***	0.626***	0.503**	0.706**
	[0.185]	[0.199]	[0.233]	[0.304]
Uncertainty x **1**(non-China)	0.072	-	-0.2	-
	[0.0998]		[0.237]	
Change in Tariff (Δln)	-4.633**	-	-13.81***	-
	[2.331]		[5.123]	
Change in Transport Costs (Δln)	-3.465***	-3.605***	-4.063***	-3.343***
	[0.240]	[0.252]	[0.447]	[0.605]
Observations	16,472	16,472	4,662	4,662
R-squared	0.15	0.19	0.24	0.38
Sector*Exporter Fixed Effects	yes	yes	yes	yes
HS6 Fixed Effects	no	yes	no	yes

Notes:
Robust standard errors in brackets clustered on HS6 industry *** p<0.01, ** p<0.05, * p<0.1 Uncertainty pre-WTO is defined as in the baseline US sample. The change in the US MFN tariff does not vary across non-preferential partners and is not identified in columns 2 and 4 when HS6 industry effects are included. Likewise, the uncertainty coefficient is not separately identified for non-Chinese imports. For columns 1-2, sample is the subset all HS6 products with imports from in 2000 and 2005 from China and one or more non-preferential MFN partner. 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.

Table 6: Differential trade effects in high vs low sunk cost industries (2000-2005)

Panel A: U.S. Import Growth (Δln) – Robustness to unobserved HS-6 import demand shocks in high vs. low sunk cost industries

	1	2	3	4	5	6
U.S. Imports From:	**China**		**Taiwan**		**Pooled**	
Sunk Cost Sample Indicator:	**Low**	**High**	**Low**	**High**	**Low**	**High**
Uncertainty pre-WTO (US)	-0.611	1.105***	-0.593	0.0486		
	[0.393]	[0.280]	[0.414]	[0.297]		
Uncertainty pre-WTO (US) x **1**(China)					0.0663	1.026***
					[0.517]	[0.383]
Change in Tariff (Δln)	-9.59	-12.09*	-6.699	-19.39***	-	-
	[13.30]	[7.107]	[11.91]	[6.484]		
Change in Transport Costs (Δln)	-2.288*	-4.512***	-3.214***	-5.181***	-1.128	-4.676***
	[1.200]	[0.743]	[0.892]	[0.724]	[0.880]	[0.797]
Observations	759	1,519	759	1,519	1,518	3,038
R squared	0.08	0.09	0.13	0.11	0.32	0.43
Sector*country Fixed Effects	yes	yes	yes	yes	yes	yes
HS-6 Fixed effects	no	no	no	no	yes	yes

Panel B: Chinese Export Growth (Δln) – Robustness to unobserved HS-6 export supply shocks in high vs low sunk cost industries

	1	2	3	4	5	6
Chinese Exports to:	**U.S.**		**EU-15**		**Pooled**	
	Low	**High**	**Low**	**High**	**Low**	**High**
Uncertainty pre-WTO (US)	0.0363	0.816***	0.492	-0.336		
	[0.313]	[0.263]	[0.302]	[0.241]		
Uncertainty pre-WTO (US) x **1**(China)					-0.444	1.144***
					[0.409]	[0.315]
Change in Tariff (Δln)	-2.659	-7.703	5.845	-12.43***	-1.15	-9.563**
	[11.49]	[5.805]	[6.075]	[3.166]	[5.881]	[3.876]
Observations	975	1945	975	1945	1,950	3,890
R squared	0.047	0.037	0.066	0.05	0.03	0.04
Sector*country Fixed Effects	yes	yes	yes	yes	yes	yes
HS-6 Fixed effects	no	no	no	no	yes	yes

Notes:

Robust standard errors in brackets clustered on HS6 industry in columns 5 and 6, *** p<0.01, ** p<0.05, * p<0.1 Uncertainty pre-WTO is defined as in the baseline US sample. Overall samples reduced relative to Tables 4 and 5 because sunk cost estimates not available for all HS-6 industries. The change in the US MFN tariff does not vary across non-preferential partners and is not identified in Panel A when HS6 industry effects are included. The uncertainty coefficient is also not separately identified for non-Chinese imports in Panel A or non-US exports in Panel B. For Panel A, the sample is the subset of HS6 products with trade in 2000 and 2005 for US imports from both Taiwan and China. For Panel B, the sample is the subset all HS6 products with trade in 2000 and 2005 exported by China to the U.S. and EU.

Table 7: Chinese Price Index Growth (2000-2005, U.S., Δln)

	1	2	3	4
Uncertainty Pre-WTO	-0.292***	-0.197**	-0.474***	-0.504**
[-]	[0.0686]	[0.0819]	[0.166]	[0.212]
MFN Tariff (Δln)	5.066***	0.585	7.920*	6.109
[+]	[1.602]	[1.678]	[4.422]	[4.627]
Transport Cost (Δln)	-0.411	-0.432*	0.733	0.801
[+]	[0.251]	[0.246]	[0.636]	[0.642]
Observations	2,579	2,579	903	903
Industry sample	HS6	HS6	HS4	HS4
R-squared	0.02	0.06	0.024	0.062
Sector Fixed Effects	no	yes	no	yes
Uncertainty Impact (Δln):[a]				
Average Price	-0.15	-0.10	-0.25	-0.26
Aggregate Price	-0.17	-0.12	-0.28	-0.30

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. Constant or sector fixed effects included but not reported. The dependent variable is the ln change in the ideal price index of Chinese varieties sold in the US between 2000 and 2005 calculated at the industry level, see the appendix for details. Sample: We use the subset of industries where value and quantity data are available and price changes are defined for at least one HS10 variety in 2000 and 2005 for the industry (HS-6 in columns 1-2 or HS-4 in columns 3-4). The HS-6 sample trims outliers in the 2.5% tails of the dependent variable.

Table 8: Chinese Variety growth (2000-2005, U.S., Δln)

	1	2	3	4
Uncertainty Pre-WTO	0.0735**	0.0605	0.245***	0.201**
[+]	[0.0309]	[0.0380]	[0.0717]	[0.0907]
MFN Tariff (Δln)	-5.178***	-4.066***	-3.196***	-2.790**
[+]	[0.957]	[1.024]	[1.179]	[1.185]
Transport cost (Δln)	-0.225**	-0.198**	-0.514***	-0.502***
[+]	[0.0913]	[0.0900]	[0.166]	[0.161]
Observations	2,579	2,579	1,051	1,051
Industry sample	HS6	HS6	HS6, uncensored	HS6, uncensored
R-squared	0.03	0.06	0.03	0.07
Sector Fixed Effects	no	yes	no	yes

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. Constant or sector effects included but not reported. The dependent variable is measured by the ln change in the number of HS-10 products in each HS6. Industry sample: same as baseline HS6 industries with price index change (columns 1, 2) and uncensored subset (columns 3, 4) that excludes those where all possible HS-10 categories were traded in both periods.

Table 9: Chinese Price Index and Variety Growth by Industry Sunk Cost Type (2000-2005, U.S., Δln)

Dependent Variable:	**Price Index**		**Variety**	
	1	2	3	4
Uncertainty Pre-WTO				
x High Sunk Cost Ind.	-0.443***	-0.296***	0.0969**	0.0828*
	[0.0822]	[0.0924]	[0.0422]	[0.0483]
x Low Sunk Cost Ind.	-0.0382	-0.0466	0.0181	0.0288
[~0]	[0.116]	[0.129]	[0.0396]	[0.0471]
MFN Tariff (Δln)	4.405***	0.386	-4.713***	-3.785***
	[1.605]	[1.684]	[0.950]	[1.012]
Transport Cost (Δln)	-0.423*	-0.437*	-0.215**	-0.194**
	[0.251]	[0.246]	[0.0900]	[0.0891]
High Sunk Cost Ind.	0.150*	0.104	0.0285	0.0286
	[0.0789]	[0.0816]	[0.0324]	[0.0343]
Constant	-0.0346		0.0551**	
	[0.0650]		[0.0230]	
Observations	2,579	2,579	2579	2,579
R-squared	0.02	0.06	0.042	0.07
Sector Fixed Effects	no	yes	no	yes

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. Dependent variables and sample described in notes to Tables 7 and 8. High Sunk cost Indicator is 1 for industries in top two terciles of export sunk cost estimates ranked by *t*-stat, as described in text.

Table 10: Chinese Export Growth (2000-2005, U.S., Δln) – Non-linear and linear estimates

	1	2	3	4
estimation method	NLLS	OLS	NLLS	OLS
Uncertainty (pre-WTO)	0.903***	0.686***	0.712**	0.567***
[+]	[0.324]	[0.150]	[0.345]	[0.185]
MFN Tariff (Δln)	-6.678***	-6.464***	-6.446***	-6.340***
[-]	[1.26]	[1.266]	[1.268]	[1.270]
Transport cost (Δln)	-4.452***	-4.309***	-4.298***	-4.227***
[-]	[0.84]	[0.844]	[0.845]	[0.847]
Constant	1.598***	0.877***		
	[0.109]	[0.0845]		
Observations	3,043	3,043	3,043	3,043
R-squared	0.02	.	0.04	.
Sector FE	no	no	yes	yes
No. coefficients estimated	3	3	23	23
Restriction test σ=3 (p-value)	0.11	n/a	0.98	n/a
Restriction test $b_\tau=b_d*\sigma/(\sigma-1)$ (p-value)	0.335	0.373	0.72	0.818

Notes
Standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.10. Columns 1 and 3 use non-linear least squares and columns 2 and 4 use ordinary least squares. Predicted sign of coefficient in brackets under variable. Sample: All specifications exclude transport cost outliers, as measured by changes in costs that exceed the top or bottom quartile by more than three times the value of the interquartile range. Uncertainty measure uses U.S. MFN (τ_1) and Column 2 tariffs (τ_2) to construct the profit loss measure. This is approximated by $1-(\tau_1/\tau_2)^\sigma$ under OLS. For NLLS we do not approximate and use instead the general function $\ln(1+b_\gamma*(\tau_1/\tau_2)^\sigma)$ where b_γ is estimated as described in the text. The four specifications in the columns restrict σ=3. We test this by relaxing the restriction in two additional NLLS specifications; we report the p-values in the 2nd to last line at which we can't reject the restriction. We also impose the restriction that $b_\tau=b_d*\sigma/(\sigma-1)$. The last row reports p-values from the test of this restriction.

Table 11: Impact of TPU reduction on Chinese export value and price (2000-2005, U.S., 100×Δln)

	Policy Uncertainty Reduction Estimates				*Data equivalent*	
	Partial Effect		General Equilibrium Effect			
	Average	Aggregate	Average	**Aggregate**	Average	**Aggregate**
Export Value	36	40	28	**32**	112	**73.0**
Export Price	-15	-17	-15	**-17**	-15.6	**-15.4**

Notes:
The first four columns are the model estimates for the change (ln) in Chinese export value (or price) due to TPU reduction. The general equilibrium employs the coefficients from NLLS estimates in table 10, column 1 and uses the model to compute and include price index effects. The partial effect estimates ignore the aggregate price index effect and use OLS estimates for exports (table 10, column 2) and prices (Table 7, column 1).The data equivalents for exports use the observed chinese export growth (Average or aggregate) and subtract nominal growth in aggregate U.S. expenditure on manufactures to account for nominal effects and aggregate expenditure shocks, which are held constant in the model prediction. The price growth data equivalent subtracts the growth in the price index of non-Chinese varieties to account for common nominal shocks. See text for additional details.

Table 12: Impact on Chinese and U.S. outcomes of TPU reduction and corresponding Tariff Ad Valorem Equivalent (2000-2005, 100×Δln)

	AVE (Tariff Equivalent)	Aggregate Change
Chinese (real) Export Value [-]	**5.0**	32.4
Chinese Export Entry & Invest. [-]	9.4	61.2
Chinese Export Price Index [+]	6.1	-16.9
US Price index (manuf) [+]	**13**	-0.5
US Consumer Welfare [-]	13	-0.4
US (real) domestic sales (manuf.) [+]	**5.9**	-1.3
US domestic employment (manuf.) [+]	7.3	-1.2
US firm Entry & Invest. (manuf) [+]	1.8	-0.4

Notes: Quantification uses NLLS estimates (column 1, table 10). AVE (advalorem equivalent) refers to the equivalent tariff reduction that would induce the same change in outcome x in the deterministic model with no uncertainty as the TPU reduction does. See text for details of calculation. Equivalents for outcome x in any period after uncertainty changes but applied policies remain unchanged. Reported values reflect differences in steady state outcomes between policy states. [-/+] denotes sign of derivative of column variables wrt tariff.

C Online Appendix (not intended for publication)

This Appendix contains of 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 General Equilibrium Model Solution

C.1.1 Algorithm

For completeness we first restate the basic elements and notation from Section 4.3 and then provide additional details on the solution algorithm and its implementation.

Basic elements and notation

- Inputs: the model and its solution require
  - A set of exogenous parameters: $\Theta \equiv \{k, \sigma, \Lambda(\tau_m, \gamma), \beta, \beta_h\}$
  - Baseline equilibrium import shares: $\mathbf{I} \equiv \{I_V(\tau_b, \gamma = 0)\}$, where $I(\tau_b, \gamma = 0) = \Sigma_V I_V(\tau_b, \gamma = 0)$.
- Equilibrium: using the entry conditions in eqs. (18) and (21) and the definitions for U, U_h we obtain a non-linear system of equations for
  - the relative stationary price index in the intermediate state: $\hat{P}_1(g, g_h, \Theta, \mathbf{I})$ in eq. (24).
  - the sequence of relative price indices after a switch to low or high protection, respectively $\hat{P}_{0,T}\left(g_h, \hat{P}_1, \Theta, \mathbf{I}\right)$, eq. (44) and $\hat{P}_{2,T}\left(g, \hat{P}_1, \Theta, \mathbf{I}\right)$, eq(45) in appendix A.2.2.
  - the average profit change due to prices after a switch to high or low protection, respectively $g\left(\hat{P}_{2,T}/\hat{P}_1, \Theta\right)$ in (20) and $g_h\left(\hat{P}_{0,T}/\hat{P}_1, \Theta\right)$ in (23).

    where $\hat{P}_{\cdot}$ denotes a price index relative to the baseline.
- Solution: $\Upsilon(\Theta, \mathbf{I}) \equiv \left\{\hat{P}_1; g; g_h; \left(\hat{P}_{2,T}; \hat{P}_{0,T}\right)_{T=0}^{\infty}\right\}$ found by
  - Fixing a set Θ consistent with our estimation and data $\mathbf{I}$.
  - Iterating n times until we obtain a fixed point such that $\Upsilon^{(n)}(\Theta, \mathbf{I}) = \Upsilon^{(n-1)}(\Theta, \mathbf{I})$.

Solution algorithm

1. Make an initial guess for $g^{(0)}$ and $g_h^{(0)}$.
2. Let $\Upsilon^{(n)}(\Theta, \mathbf{I})$ denote the values in the n-th iteration. Given two values, $g^{(n-1)}$ and $g_h^{(n-1)}$, Θ and $\mathbf{I}$ we compute the price transition paths for 250 periods $\left\{\hat{P}_{2,T}^{(n)}, \hat{P}_{0,T}^{(n)}\right\}_{T=1}^{250}$.
3. Given $\left\{\hat{P}_{2,T}^{(n)}, \hat{P}_{0,T}^{(n)}\right\}_{T=1}^{250}$ we compute updated values for $g^{(n)}$ and $g_h^{(n)}$ using

$$g^{(n)} = (1-\beta)\textstyle\sum_{T=0}^{\infty}(\beta)^T\left(\frac{P_{2T}}{P_1}\right)^{\sigma-1} \approx (1-\beta)\textstyle\sum_{T=0}^{250}(\beta)^T\left(\frac{P_{2T}}{P_1}\right)^{\sigma-1} + (\beta)^{251}\left(\frac{P_{2,250}}{P_1}\right)^{\sigma-1}$$

$$g_h^{(n)} = (1-\beta_h)\textstyle\sum_{T=0}^{\infty}(\beta_h)^T\left(\frac{P_{0T}}{P_1}\right)^{\sigma-1} \approx (1-\beta_h)\textstyle\sum_{T=0}^{250}(\beta_h)^T\left(\frac{P_{0T}}{P_1}\right)^{\sigma-1} + (\beta_h)^{251}\left(\frac{P_{0,250}}{P_1}\right)^{\sigma-1}$$

4. Check for numerical fixed point.
   - If the norm $\left\| g^{(n)} - g^{(n-1)}, g_h^{(n)} - g_h^{(n-1)} \right\| < 0.000001$, then stop.
   - Otherwise, return to step 2 using $g^{(n)}$ and $g_h^{(n)}$ as the updated starting values.
5. Check for convergence of the solution by computing the norm of difference at the steady state price index changes $\hat{P}_1^{(n)}$ and $\hat{P}_2^{(n)}$ at $g^{(n)}$ and $g_h^{(n)}$ and the terminal value of the transition price indices
   (a) To obtain the steady state solution for $\hat{P}_1^{(n)}$, we use $g^{(n)}$ and $g_h^{(n)}$ to compute U_{1V} and U_1^h and replace them in eq (24) We then directly compute $\hat{P}_2^{(n)} = \left(I_1 \hat{\tau}_2^{1-k\sigma/(\sigma-1)} + (1 - I_1) \right)^{-1/k}$.
   (b) If $\left\| \hat{P}_1^{(n)} - \hat{P}_{1,250}^{(n)}, \hat{P}_2^{(n)} - \hat{P}_{2,250}^{(n)} \right\| < 0.0001$ then stop
   (c) Otherwise: increase precision in step 4 or the number of time periods in step 3. In practice, $T = 250$ and precision in step 4 of 10^{-6} are sufficient for convergence.

Initial values and convergence

We use $g^{(0)} = \left(P_2^D / P_1^D\right)^{\sigma-1}$and $g_h^{(0)} = \left(P_0^D / P_1^D\right)^{\sigma-1}$ as the initial guess, which we compute using the deterministic equation in (16).These are upper bounds because $P_1^D < P_1^U$ and because $P_{2,T}$ and $P_{0,T}$ converge respectively to P_2^D and P_0^D from below. Using our baseline parameters and data, the algorithm typically converges to a solution in 6-20 steps for a given set of parameters. Alternative guesses, e.g. $g^{(0)} = g_h^{(0)} = 2$, take longer but converge to the same solution.

Precision and discretization

Increasing the precision beyond 10^{-6} increases computing time substantially but does not change our reported quantification results.

For our figures and quantifications over alternative values of γ or τ_1 we use 25 gridpoints. Each figure takes 2-4 minutes to produce in Matlab for Windows using a 4 core Intel processor.

C.1.2 Equilibrium Price Transition Paths

We use the multi industry version of equations (44), (45), and the definitions of U_1, U_1^h, g, and g^h to derive the price transition equations:

$$\left(\frac{\hat{P}_{0T}}{\hat{P}_1}\right)^{-k} = \frac{I_1 \sum_V r_{V1} \left(\hat{\tau}_{0V}\right)^{1-\frac{\sigma k}{\sigma-1}} + (1 - I_1) \left(\left(1 - b^{T+1}\right) + b^{T+1} \left(\frac{\hat{P}_{0T}}{\hat{P}_1}\right)^{-(k-(\sigma-1))} \left(U_1^h\right)^{k-(\sigma-1)} \right)}{I_1 \sum_V r_{V1} \left(U_{1V}\right)^{k-(\sigma-1)} + (1 - I_1) \left(U_1^h\right)^{k-(\sigma-1)}} \quad \text{for } T = 0, ... \tag{58}$$

$$\left(\frac{\hat{P}_{2T}}{\hat{P}_1}\right)^{-k} = \frac{I_1 \left(\left(1 - \beta^{T+1}\right) \sum_V r_{V1} \hat{\tau}_{2V}^{1-\frac{\sigma k}{\sigma-1}} + \beta^{T+1} \left(\frac{\hat{P}_{2T}}{\hat{P}_1}\right)^{-(k-(\sigma-1))} \sum_V r_{V1} \left(\hat{\tau}_{2V}\right)^{1-\sigma} \left(U_{1V}\right)^{k-(\sigma-1)} \right) + (1 - I_1)}{I_1 \sum_V r_{V1} \left(U_{1V}\right)^{k-(\sigma-1)} + (1 - I_1) \left(U_1^h\right)^{k-(\sigma-1)}}, \ T = 0, .. \tag{59}$$

$$U_{1V} = \left(\frac{1 + u \left(\hat{\tau}_{2V}\right)^{-\sigma} g}{1 + u}\right)^{\frac{1}{\sigma-1}}, \quad U_1^h = \left(\frac{1 + u_h g^h}{1 + u_h}\right)^{\frac{1}{\sigma-1}} \quad \text{s.t. } u_h \leq \bar{\alpha} u \ ; \ g \leq \bar{g} \ ; \ g^h \leq \bar{g}^h \tag{60}$$

We compute these using $u = \tilde{b}_\gamma / g, k = \tilde{b}_k, \sigma = 3, I_1 = .045$ and alternative $\bar{\alpha} \in \{0, 2, 4, 6\}$ as reported in Table A_params. With our estimated parameters and data alone we can compute the following weighted terms required for the multi-industry solution

$$\sum_V r_{V1}\left(\hat{\tau}_{0V}\right)^{1-\frac{\sigma k}{\sigma-1}} \tag{61}$$

$$\sum_V r_{V1}\hat{\tau}_{2V}^{1-\frac{\sigma k}{\sigma-1}} \tag{62}$$

$$\tilde{U}_1 \equiv \sum_V r_{V1}\left(1+\left(\hat{\tau}_{2V}\right)^{-\sigma}\tilde{b}_\gamma\right)^{\frac{k}{\sigma-1}-1}. \tag{63}$$

We can then replace $\sum_V r_{V1}\left(U_{1V}\right)^{k-(\sigma-1)} = \tilde{U}_1\left(1+u\right)^{1-\frac{k}{\sigma-1}}$. Similarly we compute

$$\hat{U}_1 \equiv \sum_V r_{V1}\left(\hat{\tau}_{2V}\right)^{1-\sigma}\left(1+\left(\hat{\tau}_{2V}\right)^{-\sigma}\tilde{b}_\gamma\right)^{\frac{k}{\sigma-1}-1} \tag{64}$$

and replace the term $\sum_V r_{V1}\left(\hat{\tau}_{2V}\right)^{1-\sigma}\left(U_{1V}\right)^{k-(\sigma-1)} = \hat{U}_1\left(1+u\right)^{1-\frac{k}{\sigma-1}}$.

C.1.3 Computing AVEs

As we describe in the main text, we compute AVE tariff changes that would replicate the change in outcome variables due to uncertain in our quantification. The AVE is defined as the deterministic log change in the uniform tariff factor, $\ln\Delta_y$, that generates the same change in an outcome y as TPU. Formally, Δ_y is the implicit solution to $y\left(\tau_1\Delta_y, \gamma=0\right) = y\left(\tau_1, \gamma>0\right)$. The formulas for these AVEs are in the table below in terms of $\widehat{\tau}$. We report $\ln\widehat{\tau}$ as the factor $\ln\Delta_y$ in Table 12. Note that due to the structure of the model of the implicit function for change in tariffs $\widehat{\tau}$ is the same for various outcomes, but the LHS values differ depending on the outcome variable (predicted sign in brackets).

Outcome Variable from Quantification	**Implicit Formula for $\widehat{\tau}$**
Chinese (real) Export Value [−]	$=\widehat{\tau}^{-k\sigma/(\sigma-1)}\widehat{P}(\widehat{\tau})^k$
Chinese Export Entry & Invest. [−]	$=\widehat{\tau}^{-k\sigma/(\sigma-1)}\widehat{P}(\widehat{\tau})^k$
Chinese Export Price Index [+]	$=\widehat{\tau}\left[(\widehat{\tau})^{-\frac{\sigma}{\sigma-1}}\hat{P}(\widehat{\tau})\right]^{1-\frac{k}{\sigma-1}}$
U.S. Price index [+]	$=[\widehat{P}(\widehat{\tau})]^{-1/k}$
U.S. Consumer Welfare [−]	$=[\widehat{P}(\widehat{\tau})]^{-1/k}$
U.S. (real) domestic sales (manuf.) [+]	$=[\widehat{P}(\widehat{\tau})]^{k}$
U.S. firm Entry & Invest. (manuf) [+]	$=[\widehat{P}(\widehat{\tau})]^{k}$
U.S. domestic employment (manuf.) [+]	$=[\widehat{P}(\widehat{\tau})]^{k-1}$

In practice we solve for each AVE tariff change as system of equations that satisfies the implicit functions above and a price index change $\widehat{P}(\widehat{\tau})$. Each tariff change implies a different price index, which endogenously determines exports, entry, and import price index changes. For the baseline endogenous entry model the price index change is given by

$$\widehat{P}(\widehat{\tau}) = \left[I_1\widehat{\tau}^{(1-k\sigma/(\sigma-1))} + (1-I_1)\right]^{-1/k}.$$

C.1.4 Exogenous Entry Model Solution

The exogenous entry model reference in section A.2.3 uses the same solution method, but requires fewer equations since there are not transition dynamics when applied tariffs decrease. We solve the model for g and the transition path for $\hat{P}_{2,T}$ only. Since $g^h = 1$ and therefore $U_1^h = 1$

For the AVE results, the exogenous entry model solves the same implicit formulas in the table above. The only difference is that the implicit price index change is given by

$$\widehat{P}(\widehat{\tau})^{(1-\sigma)} = I_1 \widehat{P}(\widehat{\tau})^{(k-\sigma+1)}\widehat{\tau}^{(1-k\sigma/(\sigma-1))} + (1 - I_1).$$

C.1.5 Sensitivity to Alternative α Parameters

The endogenous domestic entry model requires a value for the expected duration of an agreement, u_h, to compute general equilibrium effects. As discussed in the main text, we cannot empirically identify this parameter because the relevant domestic uncertainty factor, U_h, does not vary across industries. Our baseline parameterization assumes $\alpha \equiv u_h/u = 4$.

Our estimate of $\gamma\lambda_2 = 0.13$ implies that $\lambda_2 \in [0.13, 1]$ The range consistent with the estimates is $\alpha \in [0, 12]$. For the central case, $\alpha = 4$ we obtain $\lambda_2 = 0.28$. In Table A6, we report aggregate outcomes for exports, the share of risk in export growth, and values of λ_2 for the set $\alpha \in \{0, 2, 4, 6\}$. The export growth from reducing uncertainty is not sensitive to the choice of α, ranging from 32 to 33 log points. The share of risk is increasing in α because higher values imply lower probabilities of a bad tariff shock, λ_2. This reduces the expected mean tariff toward the current applied tariff, attributing more of the export growth to a risk reduction.

C.2 Deriving upgrading cutoffs

We focus on technology upgrades that are export market specific. If the firm has already paid the initial export entry cost, K, it can then decide to incur an additional K_z to lower its marginal export cost by a fraction $z < 1$ of the industry baseline value variable export cost unrelated to tariffs and d. Its period profits are then $\pi_v = a_s \left(zc_v\right)^{1-\sigma}$. So $z^{1-\sigma} - 1$ is the growth in period operating profits due to the upgrade. Thus, if policy is deterministic, a firm with export cost d will be indifferent between upgrading or not if its marginal cost of production is c^D_{sz}, which is defined by $\pi\left(a_s, zc^D_{sz}\right) - \pi\left(a_s, c^D_{sz}\right) = K_z\left(1-\beta\right)$

$$c^D_{sz} = \left[\frac{a_s\left(z^{1-\sigma}-1\right)}{K_z\left(1-\beta\right)}\right]^{\frac{1}{\sigma-1}} \tag{65}$$

Depending on the upgrade technology parameters we could have equilibria where the upgrading is done by all, none, or only a fraction of exporters. We focus on the latter case, which we find is the most interesting. This implies that the marginal entrant into exporting will not upgrade and therefore the entry cutoff, c^D_s, is still given by (3). Using this we can see that the upgrade cutoff is proportional to the entry cutoff by an upgrading parameter ϕ. Thus we have

$$c^D_{sz} = \phi c^D_s \tag{66}$$

$$\phi \equiv \left[\left(z^{1-\sigma}-1\right)\frac{K}{K_z}\right]^{\frac{1}{\sigma-1}} < 1 \tag{67}$$

In sum, assuming that only a fraction of exporters upgrade then the entry cutoff is unchanged and higher than the upgrade cutoff. This is assured by the restriction that $\phi < 1$, i.e. that the marginal cost reduction is sufficiently high relative to the fixed costs. Note that ϕ is independent of the policy and therefore so is the *ratio* of cutoffs.

We will now show that when only a fraction of exporters in each state upgrade then the ratio of the upgrade to the entry cutoff under uncertainty is also ϕ. This implies that the elasticity of the upgrade and entry cutoffs with respect to policy and its uncertainty are the same—a result we use in the aggregation and estimation. Given the similarities with the entry decision we will simply point out how we must modify the setup to incorporate upgrading.

We continue to assume that in any given state only a fraction of exporters upgrade so the marginal entrant in state s would not consider upgrading in *that* state. Moreover, if ϕ is sufficiently low then even the most productive marginal entrant would never upgrade, i.e. even a firm that is indifferent about entering under the worst policy state would never upgrade when conditions improved. For ease of exposition we focus on the latter case since it allows us to use the entry cutoffs derived in the main text. We will thus say that the upgrading parameter is sufficiently low if $\phi < \bar{\phi}$ and $\bar{\phi}$ is defined by $c_{0z}^{U}\left(\bar{\phi}\right) = c_{2}^{U}$ where c_{2}^{U} is the entry cutoff under column 2 tariffs previously derived and $c_{0z}^{U}\left(\phi\right)$ is the upgrade cutoff under the agreement state.

At a given state s a firm will be just indifferent between upgrading if it has cost c_{sz}^{U} , which is implicitly defined by the equality of the expected value of exporting using the upgraded technology net of the sunk cost and the expected value of waiting while using the old technology.

$$\Pi_e(a_s, c) = \max\left\{\Pi_{ez}(a_s, zc) - K_z, \beta \mathbb{E}_s \Pi_e(a'_s, c)\right\}. \tag{68}$$

The upgrade factor z multiplies the cost in the expression of operating profits for each period after upgrading. The key differences relative to the entry decision are that a firm that has not upgraded makes positive export profit today. Moreover, in the following period the firm either transitions to the same state or to column 2 tariffs, in which case it continues to use the initial technology, or transitions to the agreement state, where it will upgrade.Since z is state independent it is straightforward to show that the expected value of exporting under the new technology is simply

$$\Pi_{ez}(a_s, zc_{sz}) = z^{1-\sigma}\Pi_e(a_s, c_{sz}) \text{ for each } s \tag{69}$$

When a is decreasing in tariffs (τ) the solution is to enter when current tariffs are below a firm specific threshold tariff. We can solve the cost cutoff at any particular a by solving the upgrade indifference condition

$$\Pi_{ez}(a_s, zc_{sz}^{U}) - K_z = \Pi_e(a_s, c_{sz}^{U}) \text{ for each } s \tag{70}$$

The solution has the same form as the cutoff for investment in entry $c_{1z}^{U} = U\left(\omega, \gamma\right) c_{1z}^{D}$ and thus the relationship between the upgrade cutoff at the the entry cutoff is given by

$$\frac{c_{1z}^{U}}{c_{1z}^{U}} = \frac{c_{1z}^{D} U\left(\omega, \gamma\right)}{c_{1}^{D} U\left(\omega, \gamma\right)} = \phi$$

The proportionality of the upgrade to the entry cutoffs is analogous to the one we found under the deterministic case. Since the upgrading parameter is independent of policy values the result holds for all policy states. Moreover, the upgrade cutoff "inherits" all the properties of the entry cutoffs with respect to TPU. Namely, the upgrade cutoff under uncertainty is proportional to the deterministic cutoff in (65) by the same uncertainty factor. This also implies that the elasticity of either cutoff with respect to policy uncertainty factors is similar.

C.3 Third Country Trade Data Regressions

The data for Tables 4-6 come from the following sources. U.S. imports and transport cost measures are obtained from the NBER. We focus on non-preferential trade partners during the period 2000-2005: E.U.-15 (aggregated), Japan, Korea, Taiwan, Norway, and Switzerland. 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 comparison of China's exports to the E.U.-15 and Japan, we use reported import data at the HS-6 level obtained through COMTRADE. MFN tariff data were obtained from TRAINS. All these data are concorded to the 1996 HS revision for consistency over time. We use the same set of non-preferential trade partners in our trade participation regressions in 1996-2000 to estimate sunk export costs with the addition of Australia, which was excluded as a non-preferential partner in the 2000-2005 regressions because it implemented a PTA with the U.S. in 2005.

C.4 Double difference specification (Table A3)

If there is an industry specific growth rate trend in export growth, θ_V, and θ_V is correlated with our policy or trade cost variables, then identification is still possible via a difference-of-differences approach. Including this trend in the difference specification between 2000-2005 we have

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

where Δ_{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 Δ_{11}

$$\Delta_{11} \ln R_V = -\Delta_{11} b'_\gamma \left(1 - \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-\sigma}\right) + b_\tau \Delta_{11} \ln \tau_V + b_d \Delta_{11} \ln D_V + b' + \theta_V + u'_V. \tag{71}$$

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'_γ 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 τ_{2V} is fixed during this period and any variation in $\left(\frac{\tau_{2V}}{\tau_{1V}}\right)$ is due to small changes in τ_{1V}, already controlled for 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

$$\begin{aligned} -\Delta_{11} b'_\gamma \left(1 - \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-\sigma}\right) &\approx -\left(1 - \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-\sigma}\right) \Delta_{11} b'_\gamma \\ &= -\left(1 - \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-\sigma}\right) \frac{k - \sigma + 1}{\sigma - 1} \frac{\beta \lambda_2}{1 - \beta} (\Delta_{11} \gamma) = -\left(1 - \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-\sigma}\right) b'_\gamma. \end{aligned}$$

We then normalize each differenced RHS variable by the length of the time period to obtain magnitudes comparable to our first differenced results

$$\frac{\Delta_{11} \ln R_V}{3} = b'_\gamma \frac{\left(1 - \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-\sigma}\right)}{3} + b_\tau \left(\frac{\Delta_{11} \ln \tau_V}{3}\right) + b_d \left(\frac{\Delta_{11} \ln D_V}{3}\right) + b' + u'_V$$

This regression is similar to our OLS baseline regression in 2000-2005, but for the pre-WTO accession period 1996-1999. The main difference is that the coefficient on the uncertainty measure b'_γ reflects possibility of a change in the probability of a policy shock $\Delta_{11}\gamma$ in 1996-1999. In columns 3 and 4 of Table A3 we show this coefficient is nearly zero and insignificant. We then double difference the annualized change in exports in both periods to obtain

$$\begin{aligned} \frac{\Delta_{10} \ln R_V}{5} - \frac{\Delta_{11} \ln R_V}{3} = b_\gamma \frac{\left(1 - \left(\frac{\tau_{2V}}{\tau_{1V}}\right)^{-\sigma}\right)}{5} + b_\tau \left(\frac{\Delta_{10} \ln \tau_V}{5} - \frac{\Delta_{11} \ln \tau_V}{3}\right) \\ + b_d \left(\frac{\Delta_{10} \ln D_V}{5} - \frac{\Delta_{11} \ln D_V}{3}\right) + b - b' + u_V - u'_V \end{aligned} \tag{72}$$

The coefficients from estimating equation (72) 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 6 times the median absolute deviation from the median residuals, iterating until convergence.

C.5 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 (10) that allows the uncertainty coefficient to vary by year, t, and includes time *by* sector effects, b_{tS}, in addition to industry (HS-6) fixed effects b_V.

$$\ln R_{tV} = -b_{\gamma t}\left(1 - \left(\frac{\tau_{2V}}{\tau_{tV}}\right)^{-\sigma}\right) + b_\tau \ln \tau_{tV} + b_d \ln D_{tV} + b_{tS} + b_V + u_{tV} \quad ; t = 1996 \ldots 2006$$

We estimate two versions of this equation. First, recall that there is almost no variation over 2000-2005 in the uncertainty variable over time 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 $b_{\gamma t}^{panel} = -(b_{\gamma t} - b_{\gamma 2000}) = \frac{k-\sigma+1}{\sigma-1} \frac{\beta g \lambda_2}{1-\beta\lambda_{22}} \Delta\gamma_t$, where $\Delta\gamma_t = \gamma_{2000} - \gamma_t$. We obtain similar results to Figure A1 (from Table A4, column 1) if we drop the year 2001, constrain $b_{\gamma t}^{panel}$ to a single value for pre-WTO and a single value post-WTO (Table A4, column 2), or both. All results available upon request.

C.6 Capital intensity

We employ capital to labor intensity measures for two robustness checks. First, U.S. import growth may be higher in industries that are labor intensive and if they also have differential initial uncertainty then our estimates would be biased. In Table A5, we report the baseline estimation for the subsample in column 1 where U.S. capital intensity is from the NBER-CES Manufacturing Productivity database. In columns 2 and 3 we see that the baseline coefficient for uncertainty is not sensitive to controlling for capital intensity, nor are the coefficients for other variables. This is perhaps not surprising since we have shown that the results are robust to controlling for any type of demand (and supply shock) at the HS-6 level. Moreover, capital intensity is not significant when controlling for sector dummies, which we include in nearly all robustness checks discussed above.

Second, we check for heterogeneity in the effect of uncertainty by interacting it with the capital intensity measure. We de-mean capital intensity within the sample before interacting so that the coefficient on the uncertainty measure can be interpreted as the marginal effect at the mean capital intensity across industries. Including the control and interaction for capital intensity does not significantly affect the baseline results. Recall the model predicts a stronger effect of uncertainty for industries with export sunk costs, as we verified above. If U.S. capital intensity was perfectly correlated with export sunk costs then we should find a similar result here. In columns 4 and 5, there is a stronger effect for these industries that is marginally significant at best. Rather than suggesting some inconsistency with the model, U.S. industry capital intensity may be a poor proxy for Chinese export sunk costs at this level of disaggregation; its rank correlation with our sunk cost measure is only 0.08 in the estimation sample.

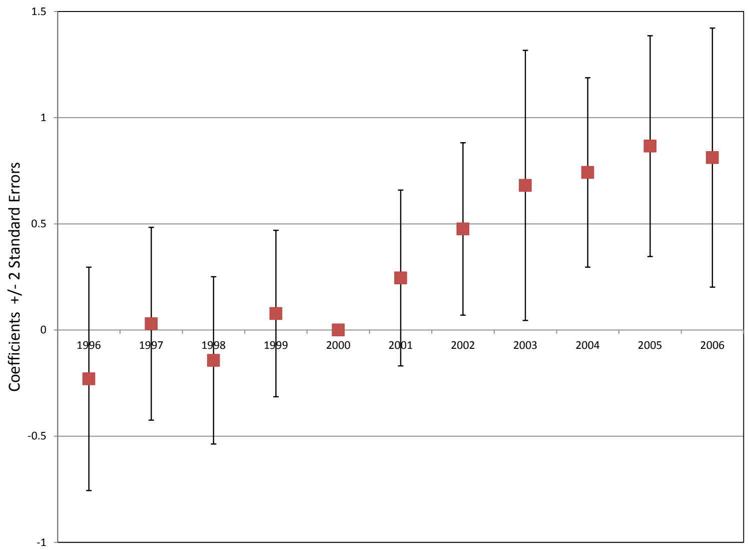
To interpret the possible explanations for these results suppose that we extended the model to include capital and that the sunk costs of production are higher in capital intensive industries. Then the weaker effects we find for the interaction of TPU with industries with higher capital intensity (relative to the interaction with high export sunk cost industries) could be due to at least two reasons. First, TPU affects export entry for incumbent producers but not enough to induce new producers that eventually export (which is exactly what the model assumes). Second, the capital intensity in the U.S. is a poor measure of Chinese investment irreversibility in production.

C.7 Capital intensity data

We concord 6 digit NAICS manufacturing codes to the 6 digit level of the HS using the correspondence in the NBER trade data. Where multiple NAICS codes match to a single 6 digit HS, we take the mean of the log of the K/L ratio. Results are robust to taking the median K/L ratio as well. Capital is measured is real dollars and labor is measured in total employment.

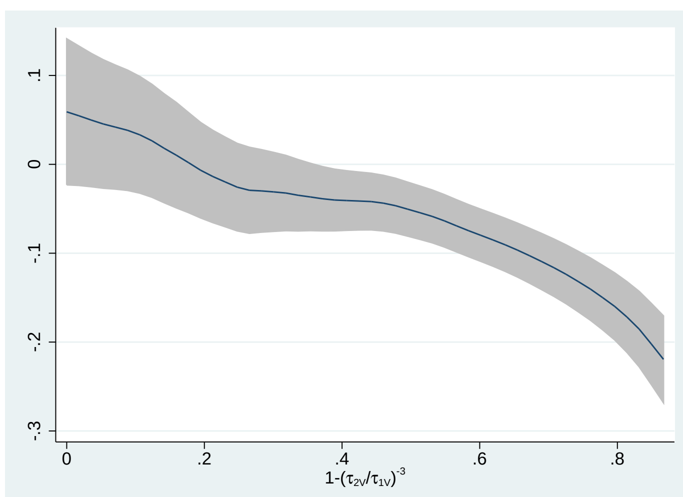
Online Appendix Figures and Tables (not intended for publication)

Figure A1: Panel Coefficients on Uncertainty Measure by Year



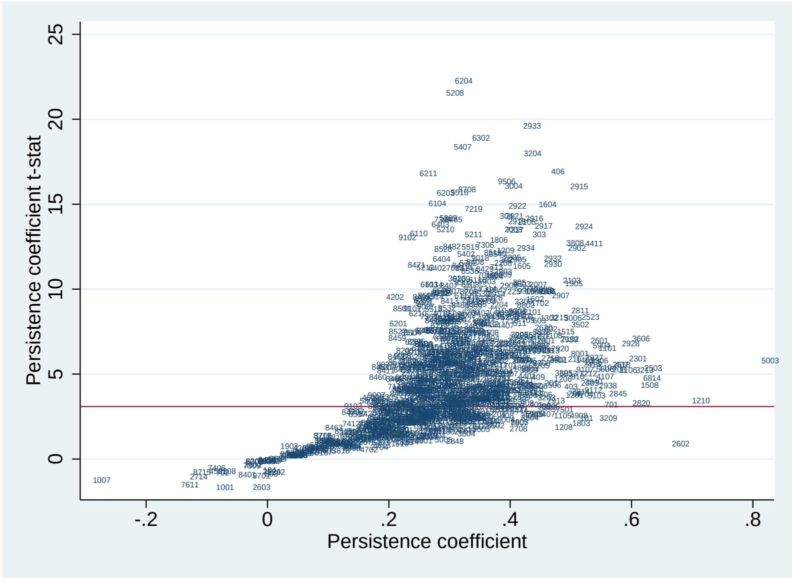
Notes: Results from an OLS unblanced panel regression on log trade flows. Uncertainty measure in 2000 interacted by year. Coefficients are changes relative to the omitted year 2000. Controls for applied tariffs, transport costs and dummy variables for section×year and HS-6 industry. Standard errors are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

Figure A2: Chinese price index ($\Delta \ln$) of continuing varieties vs initial policy uncertainty



Notes: Local polynomial fit on $1 - (\tau_{2V}/\tau_{1V})^{-3}$ where τ_{2V} and τ_{1V} are the column 2 and MFN tariff factors in 2000.

Figure A3: Sunk Cost Estimates — t-statistics vs. estimated coefficients



Notes: Estimated coefficients and t-statistics from product level persistence regressions at the HS-4 industry level described in Appendix B.3. Points are represented by the 4 digit industry code. Red line represents a t-statistic of 3.09. Two thirds of the associated t-stats are above this level.

Table A1: Uncertainty and Export Growth (Δln) by Sector — Summary Statistics

Sector	Import Share (2005)	Mean		Median		St. Dev.		C.V.		Min		Max		Obs.
		Uncertainty	Exports	Uncertainty	Exports	Uncertainty	Exports	Uncertainty	Exports	Uncertainty	Exports	Uncertainty	Exports	
1 Animals	0.26	0.15	0.43	0.00	0.68	0.18	1.42	1.27	3.29	0.00	-3.15	0.49	3.01	48
2 Vegetables	0.18	0.23	0.81	0.20	0.72	0.22	1.34	0.98	1.66	0.00	-2.28	0.60	5.20	81
3 Fats & Oils	0.00	0.25	0.55	0.29	0.58	0.18	1.21	0.71	2.21	0.00	-1.30	0.48	2.54	9
4 Prepared Foodstuffs	0.44	0.37	1.39	0.37	1.11	0.15	1.63	0.41	1.17	0.00	-2.21	0.65	7.22	68
5 Minerals	0.30	0.20	0.91	0.00	0.76	0.26	1.93	1.34	2.12	0.00	-3.96	0.73	4.66	50
6 Chemicals	1.73	0.41	1.17	0.43	1.03	0.19	1.63	0.46	1.39	0.00	-4.35	0.88	6.67	417
7 Plastics, Rubber & Articles	3.78	0.52	1.73	0.52	1.51	0.18	1.77	0.36	1.02	0.00	-3.50	0.81	7.38	148
8 Hides, Leather, & Articles	2.48	0.46	0.88	0.49	0.66	0.22	1.35	0.47	1.53	0.00	-1.05	0.77	5.73	43
9 Wood, Straw & Articles	0.89	0.46	1.95	0.54	1.39	0.20	1.66	0.45	0.85	0.00	-0.72	0.81	6.52	42
10 Pulp, Paper & Articles	1.17	0.44	1.26	0.52	1.09	0.20	1.63	0.45	1.30	0.00	-3.09	0.64	6.60	78
11 Textiles & Articles	9.73	0.64	1.50	0.70	1.24	0.17	1.87	0.27	1.25	0.00	-3.33	0.86	8.91	610
12 Footwear, Headgear, other	6.58	0.54	0.60	0.54	0.46	0.15	0.76	0.27	1.27	0.23	-1.52	0.82	2.50	53
13 Stone, Plaster, Cement, other	1.64	0.60	1.37	0.64	1.24	0.14	1.61	0.23	1.18	0.00	-3.66	0.83	7.97	123
14 Precious stones, Metals, Jewellery,...	0.94	0.48	0.94	0.50	0.86	0.31	1.45	0.64	1.54	0.00	-2.41	0.86	4.66	28
15 Base Metals & Articles	5.72	0.51	1.38	0.56	1.26	0.18	1.83	0.35	1.33	0.00	-7.59	0.84	9.97	373
16 Machinery; Elec. Equip.; Electronics	42.71	0.56	1.33	0.59	1.16	0.11	1.55	0.20	1.17	0.00	-4.36	0.83	6.86	644
17 Vehicles, Aircraft, Vessels	2.11	0.42	1.54	0.45	1.60	0.18	1.86	0.42	1.20	0.00	-4.26	0.70	7.10	81
18 Optical, Medical & other instruments	2.10	0.61	0.94	0.62	0.85	0.10	1.60	0.16	1.70	0.00	-5.20	0.86	7.80	182
19 Arms and Ammunition	0.03	0.59	1.62	0.67	1.54	0.24	1.59	0.41	0.98	0.00	-1.15	0.76	4.43	8
20 Miscellaneous Manufactures	17.09	0.65	0.86	0.65	0.73	0.11	0.97	0.17	1.13	0.25	-1.75	0.87	4.99	118
21 Art and Antiques	0.12	0.00	0.77	0.00	0.73	0.00	0.51	n/a	0.66	0.00	-0.02	0.00	1.47	7
Overall	**100.00**	**0.52**	**1.28**	**0.57**	**1.10**	**0.20**	**1.68**	**0.39**	**1.31**	**0.00**	**-7.59**	**0.88**	**9.97**	**3211**

Notes: Uncertainty measure is pre-WTO level in 2000. Exports are in log changes. Sectors correspond to the UN defined "Sections", which are coherent groups of HS-2 industries, as described in http://unstats.un.org/unsd/tradekb/Knowledgebase/HS-Classification-by-Section

Table A2: Export Growth from China (2000-2005) Robustness
Panel A: Elasticity of substitution

Potential Issue	Magnitude of common σ						Industry variation in σ	
Estimation	OLS , σ=3		OLS , σ=2		OLS ,σ=4		OLS , σ_V	
Sample change vs. baseline	none		none		none		drop V if unavailable σ_V	
			1	2	3	4	5	6
Uncertainty Pre-WTO	0.743***	0.716***	0.869***	0.838***	0.692***	0.666***	0.645***	0.592***
[+]	[0.154]	[0.186]	[0.189]	[0.226]	[0.138]	[0.168]	[0.138]	[0.149]
(…)								
Observations	3211	3211	3211	3211	3211	3211	2,733	2,733
R-squared	0.033	0.053	0.032	0.053	0.033	0.053	0.03	0.055
Sector fixed effects	no	yes	no	yes	no	yes	no	yes
Restriction p-value (F-test)	0.204	0.536	0.339	0.717	0.166	0.484	0.14	0.244

Panels B: Outliers, Selection, Specific tariffs, Processing Trade

Potential Issue			Outliers		Selection (ln growth)		Specific Tariffs		Processing Trade	
Estimation	OLS		Robust regression		OLS (midpoint growth)		OLS (AVE tariffs)		OLS	
Sample change vs. baseline	none		none		+ Rt>0, t=0 or 1		+ AVE		-(HS84x, HS85x)	
			1	2	3	4	5	6	7	8
Uncertainty Pre-WTO	0.743***	0.716***	0.521***	0.510***	0.430***	0.434***	0.956***	0.877***	0.787***	0.808***
[+]	[0.154]	[0.186]	[0.124]	[0.149]	[0.0900]	[0.108]	[0.131]	[0.154]	[0.162]	[0.198]
(…)										
Observations	3211	3211	3211	3211	3848	3841	3565	3565	2567	2567
R-squared	0.033	0.053	0.041	0.065	0.018	0.039	0.036	0.06	0.034	0.058
Sector fixed effects	no	yes	no	yes	no	yes	no	yes	no	yes
Restriction p-value (F-test)	0.204	0.536	0.002	0.021	0.0314	0.200	0.041	0.033	0.165	0.517

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 $b_d=-2.5$ for transport cost and $b_\tau=b_d(\sigma/(\sigma-1))$ can't be rejected at p-values listed in last row. Uncertainty similar to Table 2 with σ=3 except in Panel A columns 5 and 6 (uses listed values, σ_V = median estimate within HS6) and Panel B columns 5 and 6.

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.

Panel B columns 5 and 6 : use both ad valorem tariff and the ad valorem equivalent of specific tariffs (AVE=specific tariff / unit value).

Panel B: columns 7 and 8 drop HS Section XVI: machinery and electrical applicances; electrical equipment; parts thereof; sound recorders and reproducers, television image and sound recorders and reproducers, and poarts and accessories of such articles.

Table A3: Export growth from China: Robustness to HS-6 level and Pre-Accession Trends

	1	2	3	4
Dependent variable (ln):	Annualized Difference in Export Growth (2005-2000)/5-(1999-1996)/3		Pre-Accession Export Growth (1999-1996)	
Uncertainty Pre-WTO (2000)	0.506**	0.415*		
[+]	[0.224]	[0.225]		
Uncertainty Pre-WTO (1996)			-0.00501	0.0303
[~0]			[0.109]	[0.109]
Change in Tariff (Δln) [1]	-5.699***	-5.157***	-4.506***	-4.311***
[-]	[1.954]	[1.960]	[1.594]	[1.587]
Change in Transport Cost (Δln) [1]	-3.354***	-3.424***	-3.437***	-3.444***
[-]	[0.309]	[0.308]	[0.290]	[0.289]
Change in MFA quota status [1]		-0.408***		0.469***
		[0.112]		[0.160]
Change in NTB status [1]		-0.23		-0.510*
		[0.219]		[0.302]
Observations	2,571	2,571	2,571	2,571
R-squared	0.047	0.054	0.055	0.06

Notes:

Standard errors in brackets. *** p<0.01, ** p<0.05, * p<0.10. Predicted sign of coefficient in brackets under variable. Subsample of baseline observations with exports in 1999 and 1996. Robust regression employed to address potential outliers or influential individual observations due to double differencing. The estimation routine downweights outliers more than 6 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 NTB variables.

Table A4: Export Growth from China – Yearly Panel Fixed Effects Estimates (1996-2006)

	1	2
Tariff (ln)	-5.563***	-8.223***
[-]	[1.941]	[2.024]
Transport Costs (ln)	-2.468***	-2.471***
[-]	[0.226]	[0.226]
Uncertainty Pre-effect (1996-2001)		-2.179**
[-]		[0.957]
Uncertainty Post-effect (2002-2006)		-1.491
[~0]		[0.953]
Uncertainty effect relative to 2000		
1996	-0.23	
[~0]	[0.263]	
1997	0.0295	
[~0]	[0.227]	
1998	-0.143	
[~0]	[0.197]	
1999	0.0776	
[~0]	[0.196]	
2001	0.245	
[~0]	[0.207]	
2002	0.476**	
[+]	[0.203]	
2003	0.681**	
[+]	[0.318]	
2004	0.742***	
[+]	[0.223]	
2005	0.866***	
[+]	[0.260]	
2006	0.812***	
[+]	[0.305]	
Observations	37,002	37,002
R-squared	0.87	0.87
HS6 & Section by year FE	yes	yes
Restriction p-value (F-test)	0.006	0.046

Notes:

Robust standard errors with two-way clustering on HS6 and section-year, 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 σ=3. All specifications employ OLS. In column 1, uncertainty measure is fixed at 2000 level and interacted with year indicators (omitting 2000). Observations are

Table A5: Export Growth Robustness to Industry Variation in Capital Intensity

Specification:	Baseline Subsample	+ Capital Intensity Controls		+ Uncertainty Interaction	
	1	2	3	4	5
Uncertainty pre-WTO (US)	0.564***	0.706***	0.659***	0.655***	0.624***
[+]	[0.173]	[0.191]	[0.205]	[0.188]	[0.201]
Change in importer MFN tariff	-4.998***	-4.945***	-4.865***	-4.944***	-4.854***
[-]	[0.697]	[0.695]	[0.696]	[0.696]	[0.697]
Change in bilateral transport cost	-3.332***	-3.297***	-3.243***	-3.296***	-3.236***
[-]	[0.465]	[0.463]	[0.464]	[0.464]	[0.465]
Capital Intensity (K/L) in 2000 (ln)		0.0744*	0.0772	-0.0759	-0.0276
[+/-]		[0.0383]	[0.0534]	[0.0975]	[0.115]
Unc*Cap. Intensity (demeaned)				0.287*	0.193
[+/-]				[0.171]	[0.186]
Observations	3,055	3,055	3,055	3,055	3,055
R-squared	-	-	-	-	-
Sector Fixed Effects	no	no	yes	no	yes
Restriction p-value (F-test)	0.31	0.38	0.83	0.35	0.83

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 impose theoretical constraint on tariffs and transport cost coefficients: $b\tau=bd(\sigma/(\sigma-1))$. Column 1 is the baseline specification the subsample where K/L ratio is observed in the NBER-CES productivity database. Uncertainty measure uses U.S. MFN and Column 2 Tariffs to construct profit loss measure at $\sigma=3$.

Table A6: US import price growth (2000-2005, Δln) — Robustness to unobserved import demand shocks

	Matched Sample of China and non-Preferential MFN partners import price index changes		Matched Sample of China and Taiwan import price index changes	
	1	2	3	4
Uncertainty x **1**(China)	-0.274***	-0.205**	-0.226**	-0.432***
	[0.0825]	[0.0938]	[0.0901]	[0.149]
Uncertainty x **1**(non-China)	-0.0732	-	0.208*	-
	[0.0528]		[0.117]	
Change in Tariff (Δln)	-0.46	-	0.203	-
	[0.990]		[1.535]	
Change in Transport Costs (Δln)	-0.505*	-0.327	-0.282	-0.156
	[0.259]	[0.276]	[0.237]	[0.321]
Observations	4,870	4,870	3,356	3,356
R-squared	0.11	0.15	0.09	0.07
Sector*Exporter Fixed Effects	yes	yes	yes	yes
HS6 Fixed Effects	no	yes	no	yes

Notes:
Robust standard errors in brackets clustered on HS6 industry *** p<0.01, ** p<0.05, * p<0.1 Uncertainty pre-WTO is defined as in the the baseline US sample. The change in the US MFN tariff does not vary across non-preferential partners and is not identified in columns 2 and 4 when HS6 industry effects are included. The uncertainty coefficient is also not separately identified for non-Chinese imports. For columns 1-2, sample is the subset all HS6 industries with at least one continuer HS-10 variety import from in 2000 and 2005 from China and one or more non-preferential MFN partner. For columns 3-4, sample is the subset of HS6 industries continuer HS-10 traded varieity in 2000 and 2005 for US imports from both Taiwan and China. The price index dependent variable is trimmed for outliers at the 2.5% tails of the matched sample.

Table A7: Sensitivity of quantification to alternative parameterization of $\alpha = U_h/U$

α =	0	2	4	6
Implied λ_2	1	0.44	0.28	0.21
Export growth from lower uncertainty (Δln)	33.2	32.6	32.4	32.3
Growth share from Risk Reducation at Mean	0	0.54	0.71	0.78

Notes:
Each column uses a different value for α to compute the GE quantification. We use the NLLS estimates (column 1, Table 10) and include adjustments for price index effects. The share from risk reduction is the growth in exports when uncertainy is reduced from an initial equilibrium where tariffs are at their long run mean. See text for formulas.

Table A8: Summary Statistics Across Regression Specifications and Dependent Variables

Dependent Variable:	Chinese Export Growth to U.S. (ln)		Export Price Index Growth (ln)	Product Variety Growth (ln)
Table:	1-3	10	7,9	8
Change in Dep. Var. (Δln, 2005-2000)	1.28	1.26	-0.11	0.32
	[1.675]	[1.613]	[0.694]	[0.432]
Uncertainty Pre-WTO (2000)	0.52	0.52	0.52	0.52
	[0.203]	[0.199]	[0.201]	[0.195]
MFN Tariff (Δln)	-0.003	-0.003	-0.004	-0.006
	[0.00882]	[0.009]	[0.009]	[0.012]
Transport Costs (Δln)	-0.01	-0.001	-0.005	-0.007
	[0.0871]	[0.041]	[0.087]	[0.085]
Change in MFA quota status (binary)	-0.13	n/a	n/a	n/a
	[0.336]			
Change in NTB status (binary)	0.01	n/a	n/a	n/a
	[0.125]			
Observations	3,211	3,043	2,579	1,051
Estimation Method	OLS	NLLS	OLS	OLS
Fraction of total export growth	0.977	0.974	n/a	n/a

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. Product variety growth measures the number of traded HS-10 varieties within an HS-6 industry. Fraction of total export growth is the share of total export growth explained by the observations in the estimation subsample.

Table A9: Parameter Values for Quantification and Counterfactuals

Parameter	Value	Definition/Source
		Data-based inputs and assumptions for aggregate trade, price and welfare effects:
$1-\beta$	0.15	Death rate of foreign exporters
$1-\beta_h$	0.1	Death rate of U.S. firms
I_t	0.045	Chinese import penetration in 2005 to compute price effects, range is [.022, .067] from 2000-10
σ	3	Median elasticity from Broda and Weinstein (2006), rounded.
τ_0	1.038	Simple mean of MFN tariff in 2005, used to compute Figures 5-8.
τ_1	1.041	Simple mean of MFN tariff in 2000, used to compute Figures 5-8.
τ_2	1.38	Simple mean of Col. 2 tariff in 2000, used to compute Figures 5-8
		Model-based estimates for key structural quantities
k	4.45	Pareto shape parameter, Table 10, column 1, NLLS estimates
$u = \frac{\beta\lambda_{12}}{1-\beta}$	0.73	Expected spell at $m = 2$ for exporter at $m = 1$, NLLS estimate
$\alpha = u_h/u$	4	Implies U.S. firm expected to spend 4 times as long under the WTO state than a chinese exporter expected to spend under column 2.
$\hat{g}$	1.004	Computed average price effect adjustment to expected export profits for 2000
$\hat{g}_h$	0.989	Computed average price effect adjustment to expected domestic profits for 2000
		Baseline assumption for for relative spells
$\alpha = u_h/u$	4	Implies U.S. firm expected to spend 4 times as long under the WTO state than a chinese exporter expected to spend under column 2. Relaxed in Table A7 and risk quantification.

Notes: See data appendix for data sources and online appendix for exact definitions of expressions used in quantification. Estimated $\hat{g}$ and $\hat{g}_h$ are the solution implied by empirical point estimates only. The solutions are determined endogenously for each counterfactual exercises with variation over import penetration, initial tariffs, or policy shock arrival rates.

Notation Reference

Symbol	Description	Section
μ	share of income spent on differentiated goods	2.1
Ω	set of available differentiated goods	2.1
E	total expenditure on differentiated goods	2.1
p_v	consumer price of variety v	2.1
P_s	price index for differentiated goods in state s	2.1
c_v	unit labor cost for producer of variety v, the inverse of productivity $(1/c_v)$	2.1
w_e	wage in exporting country e	2.1
d_V	advalorem transport cost for industry V	2.1
$\pi(a_s, c_v)$	operating profits	2.1
K, K_z	sunk cost to start exporting or upgrading (z)	2.2,2.5
a_{sV}	demand conditions for industry V in state s	2.2
β	probability that the exporting firm survives	2.2
Π_e, Π	expected value function of exporting (e), and firm value function Π	2.2
τ_m	trade policy state $m \in 0, 1, 2$ where $\tau_2 > \tau_0$ and $\tau_1 \in [\tau_0, \tau_2]$	2.3
$U(\omega, \gamma)$	Uncertainty factor affecting entry and upgrade cutoffs	2.4
γ	policy uncertainty parameter, $\gamma \equiv 1 - \lambda_{11}$	2.4
ω	Operating profit change at col. 2 (τ_2) vs. MFN (τ_1), partial equil.: $\omega \equiv (\tau_2/\tau_1)^{-\sigma}$	2.4
$u(\gamma)$	average spell a firm starting at $s = 1$ expects to spend in state 2	2.4
λ_2	probability of state $s = 2$ conditional on exiting MFN state	2.4
ℓ	labor endowment	2.5
N_V	mass of entrepreneurs in industry V	2.5
R_{sV}	export level of industry V in state s	2.5
k	shape parameter of the Pareto distribution for productivity $G_V(c)$	2.5
α_V	industry specific distribution factor $\alpha_V \equiv \frac{N_V \sigma}{c_V^k} \frac{k}{k-\sigma+1}$	2.5
$\tilde{\alpha}_V$	industry modified factor in the export revenue $\tilde{\alpha}_V \equiv \alpha_V \left(\frac{1}{(1-\beta)K_V}\right)^{\frac{k-\sigma+1}{\sigma-1}}$	2.5
ζ_V	upgrading factor in exports for industry V, $\zeta_V \equiv 1 + \frac{K_z}{K} (\phi_V)^k > 1.$	2.5
$f(\frac{\tau_{2V}}{\tau_{1V}}, \gamma)$	general functional form for effect of uncertainty term on exports for industry V	3.1
$P_{sV,x}$	consumer import price index for industry V	3.4
$\tilde{\mu}$	parameters of indirect utility function: $\tilde{\mu} = w_e \ell \mu^\mu (1-\mu)^{(1-\mu)}$	4.1
$\hat{P}_s$	ratio of price index in state s to its baseline value	4.2
I_m	tariff inclusive import penetration in total expenditure	4.2
T	time elapsed since transition from $s = 1$	4.2
g, g_h	general equilibrium adjustment factor for average change in exporter or domestic(h) profits after a transition to high or low protection, respectively	4.2
$\hat{P}_{m,T}$	change in price index T periods after transition to tariff state m	4.3
$\alpha = u_h/u$	ratio of a domestic firm's expected spell in an agreement state $(m = 0)$ to an exporting firm's expected spell in the high protection state $(m = 2)$	5
$\lambda_{ss'}$	transition probability from state s to s' of transition matrix M	A.1