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Trade-Policy Dynamics

Evidence from 60 Years of U.S.-China Trade

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Abstract

This paper studies the growth of Chinese imports into the United States from autarky during 1950–1970 to about 15 percent of overall imports in 2008, taking advantage of the rich heterogeneity in trade policy and trade growth across products during this period. Central to the analysis is an accounting for the dynamics of trade, trade policy, and trade-policy expectations. The analysis isolates the lagged effects of past reforms and the current effects of uncertainty about future reforms. It builds a multi-industry, heterogeneous-firm model with a dynamic export participation decision to estimate a path of trade-policy expectations. The findings show that being granted Normal Trade Relations (NTR) status in 1980 was largely a surprise and that, in the early stages, this reform had a high probability of being reversed. The likelihood of reversal dropped considerably during the mid-1980s, and, despite China's accession to the World Trade Organization (WTO) in 2001, changed little throughout the late 1990s and early 2000s. Thus, although uncertainty depressed trade substantially following the 1980 liberalization, much of the trade growth that followed China's WTO accession was a delayed response to previous reforms rather than a response to declining uncertainty.

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Trade-Policy Dynamics: Evidence from 60 Years of U.S.-China Trade*

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1 Introduction

International trade depends on past, present, and future trade policy. An extensive literature studies the contemporaneous relationship between trade flows and trade policy, often summarized by a trade elasticity, while largely ignoring the effects of past or potential future trade policy.[1] One recent strand of literature has focused on the transition from past reforms, while another has focused on the effect of uncertainty about future policy while assuming that the effects of past reforms have run their course. In this paper, we reconsider the interplay between the dynamics of trade policy and trade by studying the growth of China's exports to the United States from 1971 to 2008, utilizing a new methodology that disentangles the effects of previous policy from the effects of uncertainty over future policy. Our approach simultaneously yields estimates of the dynamics of trade-policy expectations and the dynamics of the trade elasticity.

China's integration into the world trading system, and the U.S. economy in particular, is ideally suited to the study of trade-policy dynamics. It started from autarky, as the United States maintained a complete embargo on Chinese imports from 1950 until 1971. From this initial no-trade steady state, two potentially unanticipated trade liberalizations occurred in 1971 and 1980. The first granted Non-Normalized Trade Relations (NNTR) status to China, allowing Chinese exporters to enter the U.S. market for the first time, albeit at relatively high tariff rates. The second granted Normal Trade Relations (NTR) status, allowing these exporters to trade at substantially lower Most-Favored-Nation (MFN) tariffs. These two liberalizations were large, minimizing measurement concerns, and heterogeneous across sectors, providing the cross-industry variation that is key to our identification of the trade response. Moreover, beyond a limited set of goods (primarily textiles), the changes in tariffs were exogenous, having been determined outside U.S.-China trade relations; NNTR tariffs were applied to other communist countries and MFN tariffs are applied to all WTO members. The duration of these liberalizations was also uncertain, because the program to transition communist countries from NNTR to NTR rates was new in the mid-1970s

[1] The relationship between trade and trade policy is often summarized by the trade elasticity. It is generally measured in the cross section and identified by the effect of cross-country differences in current tariffs on cross-country differences in trade. Hillberry and Hummels (2013) provide an overview of the empirical estimates of trade elasticities with a limited discussion of the dynamics of the trade elasticity.

and required the United States to renew China's NTR status annually. When China joined the World Trade Organization (WTO) in 2001, the United States eliminated this renewal process, reducing uncertainty about U.S. tariffs on Chinese goods going forward.[2] The renewal process applied to all products, although the effect of non-renewal was heterogeneous across industries, owing to differences in the gap between NNTR and NTR tariffs. In total, these reforms propelled China from the United States' smallest trading partner to its largest (see Figure 1).

The key challenge in assessing the sources of China's trade integration is disentangling the effects of past policy and uncertainty about future policy. The low initial trade volumes, the scale of the reforms, and the uncertainty about the persistence of the reforms likely created a long transition that may have lasted the entire sample period. These transition dynamics make determining the impact of expected future trade policy difficult, including during the well-studied period leading up to China's accession to the WTO. The effects of future trade policy can only be measured by viewing trade growth through the lens of a model in which firms make intertemporal decisions about whether to start or stop exporting. The observed path of tariffs alone provides very little information about agents' expectations about trade policy during this transition.

We find the transition from the initial liberalizations was long lived and that uncertainty about their duration played an important role in shaping China's export growth. However, trade-policy uncertainty (TPU) was primarily important in the years following the 1980 liberalization and was only of minor importance in the years surrounding WTO accession. We estimate that the probability of a reversal of the 1980 liberalization was an order of magnitude larger in the 1980s than in the 1990s. Our results suggest much of the trade growth that occurred following WTO accession was a delayed response to the previous liberalizations.

We study the path of China-U.S. trade integration using U.S. import data from 1974 to 2008. We show that initial conditions, measured by initial trade flows or trade policies, exert a long-lasting effect on trade flows. One way of summarizing the effect of these initial conditions is to estimate a short- and long-run trade elasticity and speed of adjustment using

[2]The imposition of tariffs on Chinese imports to the United States in January 2018—and those that followed—suggests the uncertainty was not eliminated. Exactly how much uncertainty about U.S. policy decreased after China joined the WTO is a primary focus of this paper.

an error-correction model. We find that, generally, trade adjusts slowly to changes in tariffs, but China's response to U.S. trade policy is particularly slow compared with the average exporter to the United States. Our calibrated model attributes this gradualness, in part, to the very low initial trade volumes, relatively high trade elasticities when moving from near autarky, and relatively high initial uncertainty about future policy. China's integration is not unusually slow or particularly large relative to other communist/former communist countries that started near autarky or large liberalizations such as the North American Free Trade Agreement (NAFTA).

We also reconsider the evidence that WTO access reduced tariff risk and increased trade. Following Handley and Limão (2017) and Pierce and Schott (2016), we zero in on goods in which there was a large gap between MFN tariffs and the higher NNTR tariffs that would apply if NTR status was revoked, which we refer to as the NNTR gap. Theory predicts that a high likelihood of losing NTR status would depress imports of these "high-gap" goods relative to other goods, and that a reduction in this likelihood would cause relatively fast growth in imports of high-gap goods. Like these studies, we find that imports of high-gap goods grew faster than imports of low-gap goods between the early 1990s and late 2000s, suggesting that WTO accession reduced the probability of losing NTR status. We extend this analysis to start in 1974, when China had just begun to export to the United States, and find that imports of high-gap goods were depressed substantially more during the 1970s and 1980s than during the 1990s. In fact, we find that the difference in trade between high-gap and low-gap goods was largest before the 1980 liberalization, when losing NTR status was not at risk because this status had not yet been attained. During this period, the NNTR gap measures current applied tariffs—and, of course, the size of the initial move away from autarky—rather than potential tariff increases that could materialize in the future. Thus, this empirical approach cannot disentangle the role of past or present tariffs from future tariffs. We use these estimates, however, to discipline our quantitative model.

We build a model of China and the United States to study how the timing of the reforms and expectations about future trade policy shaped the transition path from the complete embargo before 1971 to the period after WTO accession. The model is a multi-industry version of the heterogeneous-firm model with new exporter dynamics in Alessandria et al.

(2021a). Industries in the model correspond directly to the goods used in our empirical analysis: the 5-digit Standard Industrial Classification of Trade (SITC), revision 2. Firms in each industry differ in terms of productivity, as well as variable export costs which they reduce gradually through a risky investment.[3] The export entry decision and gradual reduction in export costs allow for the slow adjustment of individual and aggregate trade volumes to a change in tariffs. Thus, past policy can have effects long after its implementation.

Firms make decisions to start and stop exporting incorporating expectations about idiosyncratic shocks as well as aggregate shocks to trade policy. Non-exporters decide whether to enter the export market and exporters decide whether to continue exporting. Uncertainty about future tariffs, especially if tariffs are expected to rise, can stifle export entry and encourage exit, lowering current aggregate trade volumes. The model features two trade-policy regimes, NNTR and MFN, with tariffs taken directly from the data. The probability of switching between regimes varies over time but is common across industries. The realized path of trade policy is identical to the historical experience: The model begins in 1971 in the NNTR regime and then switches to the MFN regime in 1980. We capture the initial lifting of the embargo by assuming that there are zero exporting firms in 1970.

The model is calibrated to match the historical dynamics of U.S. imports from China across industries as well as salient facts about the exporter life cycle. Most importantly, we calibrate the probabilities of switching between trade policy regimes so that the model matches our estimated time path of the annual elasticity of trade to the NNTR gap. This allows us to recover a forward-looking measure of the expected path of U.S.-China trade policy at each moment in time. Our identification works as follows. A higher likelihood of reverting from MFN tariffs to NNTR tariffs raises the expected present value of future tariffs, which lowers the expected return to exporting and thus reduces exporter entry and survival. This effect is stronger for high-gap industries than low-gap industries, reducing imports in the former relative to the latter.

We find the annual probability of China gaining access to MFN tariff rates during the 1970s was about 25 percent. Once China gained this access in 1980, the probability of losing

[3] Drozd and Nosal (2012), Fitzgerald et al. (2016), Piveteau (2021), and Steinberg (2021) develop similar models of slow firm-level adjustment to market entry through the accumulation of customers.

it was initially high, peaking at 81 percent in 1981. This observation reflects our empirical finding that trade in high-gap goods stagnated relative to trade in other goods for several years after the 1980 reform. Starting in 1986, when China applied to join the international trade arrangement that would become the WTO, this probability began to fall rapidly as trade growth in high-gap industries began to accelerate. By the mid to late 1990s, this probability was only about 5–10 percent, on par with the estimates of Alessandria et al. (2019) but lower than the estimates by Handley and Limão (2017). Joining the WTO had a minor effect on the probability of losing access to MFN tariff rates, which is reflected in the trade data: only 10–15 percent of the overall growth in imports of products with a high tariff gap occurred in the 10-year window around WTO accession (1997–2007).

Our model allows us to ask the counterfactual question: What would the transition dynamics of U.S. imports from China have looked like if the path of trade policy had been certain, rather than uncertain? In this scenario, with perfect foresight over trade policy, aggregate trade grows faster and the elasticity of trade to the NNTR gap shrinks more quickly than in the benchmark model with uncertainty (and in the data). By 1985, in the model without uncertainty, the transition is about 90 percent complete, whereas in the benchmark model, only about one third of the trade growth has been achieved. The faster trade growth is particularly strong in high-gap industries and leads to a counterfactual elasticity of trade to the NNTR gap that is less than one third of the estimated value at the start of the 1980 liberalization. A key feature of the model without uncertainty is the counterfactually high trade growth in high-gap industries in anticipation of the 1980 reform. By contrast, in the data, the coefficient on the gap grows larger during the late 1970s, consistent with a low probability of moving from NNTR to NTR during this period and a high initial probability of returning to NNTR from MFN in the early 1980s.

Our calibrated model matches the heterogeneous paths of integration across Chinese industries into the U.S. market. We use the model to recover the long-run trade elasticity from moving between steady states with different tariffs. We find that this elasticity ranges between 15 and 18 and that it increases in the initial tariff rate. These estimates are larger than those reported in the literature[4] and more than twice the elasticity of 6.5 that we

[4]Head and Ries (2001) estimate elasticities between 7.9 and 11.4 for the Canadian-U.S. FTA. Romalis

estimate from cross-sectional data. The pooled cross section mixes the short-run and long-run responses to tariffs and also contains very little tariff variation, because most of the countries besides China in our sample face the same MFN rates. Our error-correction specification distinguishes between the short run and long run and mitigates some of this limited variation. It yields long-run elasticities of 8–10, which are larger than the cross-sectional estimates but still below those implied by the model. The model is calibrated to reproduce the error-correction estimates, so the difference between these estimates and the model's true long-run elasticity implies uncertainty over trade policy makes recovering the model's true elasticity difficult.

This paper is related to the literature that studies the credibility of trade reforms. Early work focuses on the aggregate effects of temporary reforms (Calvo, 1987) or the credibility of reforms (Staiger and Tabellini, 1987; McLaren, 1997). More recent work uses firm-level and industry-level data to identify the effects of trade-policy uncertainty. This literature largely focuses on U.S. trade policy toward China. Handley and Limão (2017) and Pierce and Schott (2016) leverage the well-defined nature of the uncertainty—the annual vote to retain China's NTR status and the gap between the NTR tariffs and the fallback NNTR tariff—to measure the growth in trade that resulted from the elimination of uncertainty over future tariffs. Unlike these papers, we consider how these tariff gaps, which also capture the 1980 liberalization, influence trade dynamics in the lead-up to China receiving Permanent Normal Trade Relations (PNTR) status upon joining the WTO. Our paper also relates to several papers that use models to estimate the path of trade policy.[5] Handley and Limão (2017) also use a dynamic exporting model and the evidence on a differential change in trade to estimate the probability of NNTR reversal. Unlike their work, we estimate a time-varying probability of U.S.-China tariffs over a longer interval. Complementary to our approach, Alessandria et al. (2019) estimate a time-varying probability of a change in U.S. trade policy from 1990 to 2005, but use within-year variation in trade flows and trade-policy risk.

(2007) estimates a median trade elasticity from NAFTA of 6.9. Simonovska and Waugh (2014) estimate a trade elasticity of close to 4. Boehm et al. (2020) estimate a trade elasticity that ranges from 0.75 in the short run to about 2 in the long run.

[5] Alessandria et al. (2017) estimate the expected path of inward and outward trade policy for China from macroeconomic time series. Ruhl (2011) estimates the probability of ending the ban on Canadian beef after an outbreak of "mad cow disease."

In section 2, we provide some background on U.S.-China trade policy and trade integration. In section 3, we estimate the impact of trade policy on trade using both cross-sectional and panel methods. In section 4, we lay out the model, discuss its calibration, and present our model-implied probabilities of tariff increases. Section 5 concludes.

2 Background on Integration

Much of the focus on Chinese access to U.S. markets centers on the period leading up to China joining the WTO in December of 2001. A much larger trade liberalization, however, took place 30 years earlier. The United States embargoed trade with China from 1949 to 1971, and from 1971 to 1979, Chinese imports to the United States were subject to highly restrictive Non-Nornalized Trade Relations (NNTR) tariff rates (also known as "Column Two" rates in the U.S. Harmonized Tariff Schedule). In July 1979, diplomatic relations were restored, and China was granted access to U.S. markets at MFN rates in February 1980. This access was conditional on satisfying conditions of the Jackson-Vanik Amendment to the Trade Act of 1974 and was part of a relatively new program to provide access to U.S. markets to non-market (communist) economies.[6] Many communist or formerly communist countries accessed the U.S. market this way, although none achieved China's levels of import penetration.

The granting of MFN status in 1980, under President Jimmy Carter, was subject to annual renewal each July by the U.S. president starting in 1981.[7] From 1990 onward, under Title IV, adopted as part of the Customs and Trade Act of 1990, Congress was given 60 days from the date of the president's renewal to disapprove it, which would cause tariffs on Chinese goods to revert to the Column Two rates. In October 2000, Congress granted China Permanent Normal Trade Relations (PNTR) upon China's agreement to join the WTO. In December 2001, China officially entered the WTO, eliminating the annual renewal of NTR

[6]China was the third country to receive a waiver, following Romania in 1975 and Hungary in 1978. After China, no country gained access for another 10 years, and in the meantime, Romania lost access from 1988 to 1991.

[7]The change in administration in 1981 likely increased uncertainty, as President Ronald Reagan adopted more protectionist policies and was more openly anti-communist.

status.[8] From 1990 to 2000, Congress only voted once against renewing China's NTR status.[9]

These policy changes effectively took tariffs on Chinese imports from infinity (1949–1970) to NNTR rates (1971–1979) and finally to MFN rates (1980–2017). Figure 2 plots the path of the distribution of tariffs, summarized by the median tariff and the interquartile range, from 1974 to 2008. The median tariff fell from 30 percent to about 8 percent. Subsequent declines in tariffs were related to phaseouts from successive rounds of the General Agreement on Tariffs and Trade (GATT; the Tokyo round in 1980–1986 and the Uruguay round in 1994–1999). To capture a broader view, including the embargo, and to control for these multilateral GATT reforms, we plot China's inverse tariff rate in Figure 4. The *inverse tariff rate* is the inverse of the Chinese tariff, on good g at time t, normalized by the NTR rate,

$$\tau_{gt}^{\text{INV}} = \frac{1 + \tau_{NTR,gt}}{1 + \tau_{CH,gt}}, \tag{1}$$

which ranges from zero, when Chinese goods were embargoed, to one, when China accessed U.S. markets at MFN rates. In the figure, we plot the median good and the interquartile range. It shows only two large changes in tariffs: they occur in 1971 and 1980. In 1971, the inverse tariff rate jumps from zero to 80 percent, and in 1980, it jumps to one, eliminating the tariff gap. Unlike the GATT rounds, which are characterized by agreed-upon tariff phaseouts, these tariff cuts were immediate. This is an advantage of investigating the Chinese experience: phaseouts make disentangling the expectations of future tariffs from current tariffs virtually impossible. The path of U.S. tariffs on Chinese goods was governed by a very different process than the path of tariffs observed in other liberalizations, such as WTO rounds, or from free trade agreements.

Figure 1 provides some perspective on the importance of trade by China with the United States. The left panel shows China exports lagged in its share of world GDP through the mid 1980s, as China was initially quite closed. Comparing China's share of global non-U.S. exports with China's share of U.S. imports in the right panel, we see the United States was relatively closed to China until the mid to late 1980s. The right panel of Figure 1 plots

[8] Additionally, access to the WTO brought China into the Agreement on Textile and Clothing that was scheduled to phase out the quota system in four steps ending in 2005.

[9] In 1992, the non-renewal resolution passed in the House and Senate but was vetoed by President George H. W. Bush.

U.S. imports from China as a share of overall U.S. imports and shows the post-1980 period featured sustained growth in U.S. imports from China, although at a slowing rate.

The entry of new firms into exporting—growth on the *extensive margin*—is an important way that aggregate trade grows in response to a tariff cut (Kehoe and Ruhl, 2013; Alessandria et al., 2020), and this is an important mechanism in our quantitative model. In Figure 3, we plot a measure of the extensive margin for exports from China to the United States. The fraction of SITC goods that China exports to the United States grows gradually from about 0.2 in 1974 to 0.6 in 2008.[10] In the median NTR country, this measure is flat.

The prolonged growth of Chinese exports could be due to the gradual response of trade to the very large past reforms, or it could be due to the expectations of future tariff reversals becoming less likely. We explore these two possibilities, next, taking advantage of the rich heterogeneity in trade policy in this period.

3 Empirical Evidence

This section documents two main facts about China's NTR tariff liberalization. First, China's trade response was especially gradual relative to the average trade response to tariff changes. Second, the level of tariffs prior to the NTR liberalization has a persistent role in explaining the variation across goods in the volume of U.S. imports from China. Consequently, the delayed effect of the NTR liberalization is instrumental in accounting for the findings of the TPU literature regarding the impact of China's PNTR status grant in 2000 (Pierce and Schott, 2016; Handley and Limão, 2017).

3.1 Data

We work with annual U.S. imports from 1974 to 2008, aggregated at the 5-digit level of the Standard International Trade Classification (SITC), revision 2.[11] We refer to this level of aggregation as a *good* and denote it by g. Our data include applied duties, cost-insurance-and-freight (CIF) charges, and the free-on-board (FOB) value of imports. The log FOB value of imports is denoted by v_{gjt}, where j indexes the exporting country and t indexes

[10] Ideally, we would like to measure the number of new firm-goods being traded, but we lack data on firm-level trade in China going back to the 1970s.

[11] The data are from Feenstra (1996). Although more disaggregated import data are available, building a consistent concordance is difficult and noisy.

time. Tariffs on imports, τ_{gjt}, are calculated as the applied duties divided by the FOB import value.

The ad-valorem-equivalent NNTR and NTR statutory rates are obtained from Feenstra et al. (2002) at the 8-digit level of the Harmonized Tariff Schedule (HS). We assign HS-8 products to 5-digit SITC rev. 2 goods using the concordance from United Nations Trade Statistics (2017), and we define the SITC-level tariff as the median tariff within the good. The *NNTR gap* is the difference between the NNTR rate and NTR rate, which can vary over time as NTR rates change. Following Pierce and Schott (2016), we use the NNTR gap in 1999 as our measure of "the" NTR gap.[12,13]

Our baseline sample of exporters to the United States includes China and all the other countries that had normal trade relations with the United States throughout the entire sample period and were not a part of preferential trade agreements (PTA) with the United States. These conditions exclude Canada and Mexico, as well as the other communist countries.[14] A key feature of our baseline sample is that after 1980, all sources should face the same NTR rates with few exceptions.[15] We also exclude all of the goods that include products that were subject to quota removals under the Multi Fiber Arrangement (MFA), as documented in Bambrilla et al. (2010). Table 1 reports the distribution of tariff rates by 1-digit SITC industries by NNTR rates and applied rates. The mean NNTR rate is 28 with a standard deviation of 18. The average applied rates are five in 1979 and two in 2001. For China, the average applied rate is considerably higher than other countries in 1979 (20 versus five), but quite similar to other countries by 2001 (three versus two). Thus, tariffs generally vary little across countries receiving NTR rates, and most of the variation in tariffs in our analysis is between NNTR and NTR rates in the early part of the sample.[16]

[12]All of our results are robust to using the mean HS-8 NNTR gap, as well as a time-varying NNTR gap.

[13]The simple correlation between the NNTR gap and the tariffs imposed on China in 2018 and 2019 at HS-8 level is -0.06. This finding suggests the NNTR gap ceased to be an effective measure of trade-policy risk and justifies our baseline sample-period end.

[14]The list of countries excluded because they held NNTR status at some point in the sample period are Albania, Bulgaria, Cambodia, Cuba, Czech Republic, Hungary, the Democratic People's Republic of Korea, Romania, the Slovak Republic, Vietnam, and the 15 countries that formed the Soviet Union.

[15]Applied tariffs would differ owing to aggregation, specific tariffs, temporary commercial policy, or measurement error.

[16]The lack of tariff variation among the NTR countries creates some concerns about the elasticities being estimated across countries receiving NTR rates for the whole sample. We return to this in section 4.5.

3.2 Gradual Adjustment to Trade Liberalizations

The 1980 liberalization is ideal for studying the dynamic adjustment to trade policy. The liberalization occurred in a single step, and China's eventual WTO accession was still 20 years away—and likely considered a very low-probability event. These conditions largely eliminate the confounding effects of phased-in liberalizations and anticipation effects on the observed export dynamics. Although other changes occurred in China during our period of study, the gradual growth in trade volumes that we observe likely reflects the gradual adjustment to the 1980 liberalization.

A convenient way to characterize the adjustment process is through short-run and long-run elasticities (Ruhl, 2008). We estimate short-run and long-run import elasticities using two variations of the error-correction method (ECM) (Johnson and Oksanen, 1977; Johnson et al., 1992; Gallaway et al., 2003). The first is the *restricted version* of the ECM, also known as the partial-adjustment model. We estimate

$$\begin{aligned} v_{jgt} =& \sigma^{SR}_{China}\mathbb{1}\{j = China\}\tau_{jgt} + \sigma^{SR}\mathbb{1}\{j \neq China\}\tau_{jgt} \\ &+ \alpha_{China}\mathbb{1}\{j = China\}v_{jg,t-1} + \alpha\mathbb{1}\{j \neq China\}v_{jg,t-1} + \delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}, \end{aligned} \tag{2}$$

where σ^{SR} is the short-run (annual) elasticity, and the long-run elasticity is $\sigma^{LR} = \sigma^{SR}/(1-\alpha)$. The lagged import term yields the speed-of-adjustment coefficient, $1/(1-\alpha)$. We also include country-year (δ_{jt}) fixed effects to capture aggregate shocks to exporting countries and good-year fixed effects (δ_{gt}) to capture good-level U.S. demand shocks. By including country-good fixed effects (δ_{jg}), estimates should be viewed as exports of jg relative to 2008.

The *unrestricted version* of the ECM is

$$\begin{aligned} \Delta v_{jgt} =& \sigma^{SR}_{China}\mathbb{1}\{j = China\}\Delta\tau_{jgt} + \sigma^{SR}\mathbb{1}\{j \neq China\}\Delta\tau_{jgt} \\ &+ \alpha_{2,China}\mathbb{1}\{j = China\}\tau_{jg,t-1} + \alpha_2\mathbb{1}\{j \neq China\}\tau_{jg,t-1} \\ &+ \alpha_{1,China}\mathbb{1}\{j = China\}v_{jg,t-1} + \alpha_1\mathbb{1}\{j \neq China\}v_{jg,t-1} \\ &+ \delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}. \end{aligned} \tag{3}$$

The addition of the lagged tariff allows for a more flexible short- to long-run adjustment. Note the dependent variable is the 1-year log difference and the short-run elasticity, σ^{SR}, is the coefficient on the 1-year log difference in tariffs. The long-run elasticity is now $\sigma^{LR} = -\alpha_2/\alpha_1$.[17] We maintain the same set of fixed effects as in equation (2).

In Table 2, we report the results of estimating equations (2) and (3) for U.S. imports from China and the other NTR countries. To obtain separate estimates, we interact each of the terms on the left-hand side of the estimating equations with an indicator variable for China. In column 1, we report the cross-sectional estimates, that is, estimates from equation (2) without the lagged import term. Not surprisingly, given China's large export growth, its average trade elasticity is 2.5 times larger than the average of the NTR countries.[18]

Column 2 reports the results of the restricted ECM. China's long-run elasticity is 7.8, almost three times larger than its short-run elasticity of 2.8. For the NTR countries, the short-run elasticity is 1.9 and the long-run elasticity is 4.0, a ratio of about two. Column 3 reports the estimates of the unrestricted ECM. There are two key takeaways. First, China's larger trade elasticity reported in column 1 is almost entirely driven by the long-run adjustment, because the annual trade elasticity is close to that of the NTR countries. Second, China's ratio of the long-run to short-run elasticity (3.5) is more than twice as large as that of the NTR countries (1.7).

Table 3 shows that our finding that trade flows adjust relatively slowly to changes in trade policy is robust to a range of alternative specifications. Columns 2 and 3 include shipping costs (CIF charges) and dummies for entry and exit of goods as additional controls. Columns 4 and 5 show our results are robust to including all countries and MFA goods in our sample. Column 6 reports results from a specification in which we only include goods that had non-zero U.S. imports from China at some point in the period 1974–1979. Finally, columns 7 and 8 illustrate that China's speed of adjustment to the NTR liberalization is comparable to that of Canada's and Mexico's in response to NAFTA, when tariffs were gradually eliminated

[17] Note one can transform equation (3) into levels; then, the long-run elasticity becomes $(\alpha_2 - \sigma^{SR})/(1-\alpha_1)$. By setting $\alpha_2 = 0$, this expression is the same as the long-run elasticity under the restricted ECM.

[18] When we restrict the sample of countries to Canada, Mexico, and the ex-communist countries, China's elasticity is only 1.7 times larger than the other-country average. These countries experienced large heterogeneous liberalizations compared with the NTR countries in the benchmark sample, suggesting the magnitude of tariff changes plays an important role.

over periods of up to 15 years. Although the delayed effect of NAFTA's tariff reductions are not necessarily due to the phaseout of tariffs (Besedes et al., 2020), these similarities suggest the adjustment to major trade liberalizations can take a long time.

3.3 The Delayed Effect of China's NTR Liberalization

We first document that the NNTR gap, commonly used in the TPU literature to capture the risk of a reversal of China's NTR status, was strongly related to the applied duties on China's exports to the United States before 1980. We then revisit the effect of the NNTR gap on China's exports to the United States after its PNTR access, using our longer sample period. Third, we show a strong trend in the effect, and it can be traced back to the 1980 NTR liberalization.

3.3.1 Revisiting the Effect of the PNTR Access

We revisit the results in Pierce and Schott (2016), changing only the sample period and the aggregation level of trade flows. Pierce and Schott (2016) focus on HS-8 product trade flows between 1992 and 2007. We use 5-digit SITC goods between 1974 and 2008 to estimate

$$v_{jgt} = \beta \mathbb{1}\{t < 2000\}\mathbb{1}\{j = China\}X_g + \sigma\tau_{jgt} + \delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}, \tag{4}$$

where X_g is the NNTR gap for good g. This equation is the same as (2), except it lacks lagged imports and includes the interaction between the China dummy and the pre-PNTR dummy. Note $\beta < 0$ indicates China's exports of high NNTR-gap goods were either relatively high in the post-PNTR period or relatively low in the pre-PNTR period. Although this specification is not able to distinguish between the two cases, the common interpretation in the TPU literature is the former.

Table 4 reports the results. Column 1 uses the same sample period as in Pierce and Schott (2016) but a coarser level of aggregation. The coefficient of interest, β, measures the effect of the NNTR gap on a good's export growth after China's PNTR accession. We find that $\beta = -0.9$, compared with -0.5 in Pierce and Schott (2016). Column 2 reports the result for our full sample period. We find that $\beta = -2.5$, a 2.5 times increase from the estimate obtained using the shorter sample period. Column 3 reports our results when we use our baseline sample of countries and the longer sample period. In this case, β increases further to -3.3.

Column 4 excludes applied tariffs from equation (4), because we exclude those from our baseline estimation of the annual effect in the next subsection. The coefficient on the pre-PNTR dummy is now $\beta = -3.6$, which is only a small increase from our benchmark specification ($\beta = -3.3$). Column 5 shows the result is similar when we use the NNTR rate instead of the NNTR gap, which is the result of the good-year fixed effects absorbing the variation in MFN rates. Finally, column 6 uses the applied NNTR rate instead of the NNTR gap in 2001. As is evident in Figure 5, the NNTR statutory tariffs (a part of X_g) are correlated with the applied tariffs (τ_g). The correlation between the mean SITC applied tariff rate over the period of 1974 to 1979 and the median NNTR rate is 0.78. The effect of the pre-PNTR dummy is even stronger when we use applied NNTR rates instead of the scheduled NNTR gap. Overall, these findings indicate the effect of NNTR rates on China's exports to the United States after its PNTR access becomes significantly larger as the sample period increases.

3.3.2 The Annual Effect of the NNTR Gap and Chinese Exports

We now explore the time path of the effect of the NNTR gap on China's exports to the United States by estimating

$$v_{jgt} = \sum_{t'=1974}^{2007} \beta_t \mathbb{1}\{t = t'\} \mathbb{1}\{j = China\} X_g + \delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}. \quad (5)$$

Note this equation is the same as equation (4), except that the effect of the NNTR gap is now interacted with a year dummy, and we exclude the tariff rate because it is highly correlated with the applied NNTR rate prior to 1980. Therefore, its inclusion would undermine the explanatory power of the gap during these early years.[19] This issue was less of a concern when we considered only the shorter time period. The MFN rate in the gap measure, however, is not an issue. The sample is restricted to NTR countries, so the good-year fixed effects control for changes in scheduled MFN rates. Finally, note the elasticities $\{\beta_t\}_{t=1974}^{2007}$ capture

[19] We report estimates of equation (5) that include tariffs in Table 3. The β_t coefficients in the earlier years are, in fact, smaller than those from the specifications that do not include tariffs.

the effect of the NNTR gap of Chinese exports relative to the export level in 2008. Hence, we expect the elasticities to be negative if growth in higher-NNTR-gap goods was depressed or still adjusting to the NTR liberalization prior to 2008.

The results of estimating equation (5) are illustrated in Figure 6 and reported in column 1 of Table 5. The trend in the effect of the NNTR gap is striking. Between 1974 and 1979, the effect is relatively stable around an elasticity of -10, indicating high-NNTR-gap goods were significantly depressed relative to low-gap goods before the NTR liberalization in 1980. After 1980, the elasticity starts to rise gradually as high-NNTR-gap goods catch up with low-NNTR-gap goods. This adjustment is slow and the coefficients level off at about -1.6 in 1993. From 1991 to 1993, the elasticity remains relatively stable between -1.7 and -1.1. The slowdown in the adjustment between 1992 and 1999 coincides with the introduction of the annual vote on China's NTR status by the U.S. Congress. After 1999, the elasticity rises above -1 and becomes statistically insignificant, suggesting the adjustment to the NTR liberalization was completed by 1999.

Figure 6 presents three main takeaways. First, the large effect of the NNTR gap on China's export growth can be traced back to the NTR liberalization. Second, the adjustment to the NTR liberalization was very persistent and likely to have been slowed by to the introduction of TPU. Third, the positive association between post-PNTR growth and the NNTR gap found in Pierce and Schott (2016) and Handley and Limão (2017) is mostly due to low pre-PNTR growth, as opposed to high post-PNTR growth.

The difference-in-differences approach summarized in Figure 6 cannot differentiate between the influences of uncertain future policy and past reform on trade volumes. The quantitative model we develop in the next section is our best tool for measuring these two influences, but a simple modification to equation (5) suggests that the lagged effect of past policy may be important. In Figure 7, we report the estimates of a version of equation (5) in which we have added lagged trade volumes, $v_{jg,t-1}$, to the list of regressors. This model allows for trade volumes to be serially correlated, which we would expect if aggregate trade takes time to adjust to a cut in tariffs. Conditioning on lagged trade volumes decreases the elasticity of trade to the NNTR gap by about 50 percent in the 1970s and the elasticity is significantly lower than the baseline estimates until the 1990s.

3.3.3 Robustness

The time-varying pattern of the effect of the NNTR gap on China's exports to the United States, illustrated in Figure 6, is robust to the following range of alternative approaches reported in Table 5 and Table 6.

Anticipation 1979. Figure 2 shows Chinese exports to the United States were rising robustly in 1979, but our empirical framework suggests this increase was smaller for high-gap goods, because the elasticity estimate falls in 1979. The weak growth for goods whose tariffs were about to decline could reflect some delays in purchases followed by boosted purchases as emphasized by Khan and Khederlarian (2019). To control for the possibility of these anticipatory effects, we estimate equation (5) including the lead change in applied tariffs ($\Delta\tau_{jg,t+1}$). Effectively, only in 1979 is this control relevant, because in other years, changes in applied tariffs are minimal (see Figure 4). Column 2 of Table 5 shows this inclusion indeed smooths the response of trade flows to the NNTR gap around in 1979, and we no longer see a drop in the elasticity.

Additional Trade-Cost Controls. The baseline estimation of equation (5) departs from that in Pierce and Schott (2016) by excluding applied tariff rates τ_{jgt}. We excluded the tariffs to avoid problems caused by the high correlation between the NNTR rate and the applied tariffs before the liberalization. Column 3 of Table 5 includes shipping costs, calculated as CIF charges over the FOB value of imports. The coefficients of interest, $\{\hat{\beta}_t\}_{t=1974}^{2007}$, are virtually unchanged. In column 4, the specification also includes the applied tariff rates. Consistent with the correlation between applied tariffs and the NNTR gap, its inclusion biases downward the coefficients of interest in the years before the liberalization. After 1980, the coefficients of interest are almost identical to those in column 3; in the years before the liberalization, they drop slightly.

Good-Level Life-Cycle Controls. Given the long time horizon of the analysis, we study the robustness of the baseline results to the inclusion of variables that capture a good's life cycle. Namely, we introduce the following variables in equation (5): age, which we define as the number of years a good has been sourced from a country j; age squared; and dummy variables for the first, second, penultimate, and last years a good appears in the sample. Column 5 of Table 5 shows their inclusion has only a minimal impact on the baseline results.

Sample of Countries and Goods. In columns 1–3 of Table 6, we report the robustness of our results to the restrictions on countries and goods of our baseline sample. Column 1 reports results from the specification that includes all countries. Column 2 reports results from the specification that also relaxes the MFA exclusion. Whereas column 5 shows that including all countries has no impact on the baseline results, the coefficients in column 2 drop by 10–20 percent. The drop is common across all years, however, leaving the speed of adjustment and the overall pattern of the baseline unchanged. The baseline sample included goods that were not traded in 1974–1979 but were traded later. Column 3 restricts the sample to those goods in which U.S. imports from China were non-zero at some point before 1981. The coefficients of interest are almost identical to the baseline results, although, interestingly, a slightly longer effect of the NNTR gap remains after the PNTR liberalization in 2001.[20]

Alternative Measures of the NTR Liberalization. The baseline measure of the NTR liberalization considers the good-level NNTR gap in 2001 to map the results into those of Pierce and Schott (2016). Column 4 of Table 6 defines X_s to be the NNTR rate. Although the coefficients of interest, $\{\hat{\beta}_t\}_{t=1974}^{2007}$, are 10–20 percent smaller in the first years, the gradual nature of the adjustment is unchanged. Column 5 defines X_s as the applied NNTR rate, calculated as the average good-level tariff applied on imports from China between 1974 and 1980. The results with the applied NNTR rate are very similar to the baseline. Whereas $\{\hat{\beta}_t\}_{t=1974}^{2007}$ are around 5-15 percent smaller than the baseline throughout all years, they converge to zero (or statistical insignificance) slightly faster in the last five years before the PNTR liberalization.

Level of Aggregation. Our baseline level of aggregation is at the 5-digit SITC level, because we can construct a continuous dataset for the sample period of 1974–2008 and include

[20]We have also experimented with extending our sample period from 1974–2008 to include 1970–1973 and 2009–2017. Between 1970 and 1973, Chinese exports to the United States are insufficient to yield significant estimates. However, when we pool over those years, the effect is similar to that in 1974. Extending the sample period until 2017 yields an additional bump of around -1 percentage points in the elasticity of trade to the NNTR gap.

the 1980 NTR liberalization. Nonetheless, the results with more disaggregated product-level data are very similar to the baseline when we consider the periods 1974–88 (using the 8-digit TS-USA product-level aggregation) and 1990–2008 (using the HS-8 product-level aggregation), separately. Figure 8 plots the time-varying effect of the NNTR gap on exports from China obtained from equation (5) when we use the 8-digit TS-USA (Tariff Schedule of the United States) product-level aggregation for the period 1974–1987.[21] The pattern is strikingly similar to the baseline pattern with SITC aggregation. Figure 9 uses the HS-8 product-level aggregation for the period 1989–2007. Between 1989 and 1993, the negative effect is larger, rising from around -3.43 in 1990 to around -1 in 1993. After 1993, it remains relatively stable at -1. In the aftermath of China's PNTR access, the effect halves and then becomes statistically insignificant by 2002.[22] Note the summation of the total elasticity under TS-USA (-4), HS-8 (-3), and the two years under SITC (-2) yields a total elasticity of -9, which is close to the baseline estimate. More importantly, the time-varying pattern of the effect is also consistent with the baseline results. These results suggests our baseline findings are robust to the level of aggregation used.

China Supply Effects. Our baseline approach controls for U.S. demand shocks specific to particular goods but not for good-specific Chinese supply shocks. To study the importance of these potentially confounding effects, we use the World Trade dataset from Feenstra et al. (2005) (1974-2000) merged with the BACI Trade database (2001-08).[23] We estimate

$$\begin{aligned} v_{ijgt} = &\sum_{t'=1974}^{2007} \beta_t \mathbb{1}\{t = t'\}\mathbb{1}\{i = U.S.\}\mathbb{1}\{j = China\}X_g \\ &+ \mathbb{1}\{i = U.S.\}\mathbb{1}\{j = China\}X_g + \delta_{igt} + \delta_{jgt} + u_{ijgt}, \end{aligned} \tag{6}$$

where δ_{jgt} controls for exporter supply shocks and δ_{igt} for importer demand shocks (and trade

[21]The NNTR gap is not available at the TS-USA product aggregation. Instead of the NNTR gap, we use the mean TS-USA applied tariff on U.S. imports from China between 1974 and 1979 to proxy for the NNTR rate. Figure 5 shows applied NNTR rates are closely related to their schedule.

[22]Tables B.1 and B.2 in the appendix provide some robustness checks of the described pattern.

[23]Note this dataset is at the SITC 4-digit level. We include the 50 largest exporter countries in 2001 except Hong Kong SAR, China; Canada; Mexico; and the former Soviet Union members and aggregate the remaining countries into one. We further exclude goods subject to the MFA quotas as in our baseline. None of these restrictions change the results.

barriers). Note our coefficients of interest now include the difference in China's exports to the United States versus other destinations on top of the triple difference from our baseline approach. The coefficients are relative to the effect in 2008 as in the baseline by including the term $\mathbb{1}\{i = U.S.\}\mathbb{1}\{j = China\}X_g$. Figure 10 reports the estimates of $\{\beta_t\}_{t=1974}^{2007}$. The overall pattern is similar to the baseline. Indeed, we cannot say the path of the gap coefficient with China supply factors is statistically different from our baseline.

4 Quantitative Analysis

To quantify the contributions of gradual adjustment to past reforms and expectations about future policy changes to the growth of U.S. imports from China, we construct a partial-equilibrium model of export participation dynamics and calibrate it to match the empirical evidence documented in section 3.

4.1 Model

The model consists of G goods that correspond to the 5-digit SITC goods in the empirical analysis. Within each good g, a continuum of firms produce differentiated varieties. Firms are characterized by their productivity (z) and variable trade cost (ξ). Firms die exogenously at a rate of $1-\delta(z)$, where firms with higher productivity have a lower probability of death. The mass of firms in each good g is fixed: when a firm dies, it is replaced exogenously by a new firm. To export to the United States, a firm must pay a fixed cost that depends on whether it exported in the previous period. Two trade-policy regimes, NNTR and MFN, exist, and the probability of switching between regimes varies over time.

Production and demand. Firms operate constant-returns-to-scale technologies that use labor as the only input,

$$y_t = z_t \ell_t, \tag{7}$$

where z is a firm's productivity. Productivity follows a stationary first-order Markov process. U.S. demand for a firm's good, d_{gt}, is a downward-sloping function of the firm's price, p,

$$d_{gt}(p, \xi_t, \tau_{gt}) = (p\tau_{gt}\xi_t)^{-\theta} D_{gt}, \tag{8}$$

where τ_g is the current U.S. tariff on goods of type g, ξ is the firm's idiosyncratic variable trade cost, and D_g is an aggregate demand shifter that is common to all firms in good g. The firm's demand has elasticity θ, but it will not be the elasticity of aggregate trade to a change in tariffs. The aggregate trade elasticity is determined by θ and the export participation response of the tariff change.

Trade costs. Firms face two types of costs to access the U.S. market. These costs are technological—they are not policy variables. The first cost is a stochastic variable cost, ξ, which represents the efficiency with which a firm can transform a unit of goods in China into a unit of goods in the United States. This cost can take three values ($\infty > \xi_H > \xi_L$) and evolves according to a stationary, first-order Markov process. When $\xi = \infty$, the firm is a nonexporter. When ξ is finite, some firms will choose to export. Following Alessandria et al. (2021a), we assume the probability of retaining the current ξ is symmetric: $\mathrm{P}(\xi_L|\xi_L) = \mathrm{P}(\xi_H|\xi_H) = \rho_\xi$. This specification implies exporters start small and, with some luck, grow large.

The second type of trade cost is a fixed cost, f, that the firm must pay in order to export in the next period. The fixed costs are identical across firms. If the firm is currently a nonexporter, it pays f_0 to enter the export market next period. If the firm is currently exporting, it pays $f_1 \leq f_0$ to continue exporting in the next period. We summarize the fixed-cost structure in a function, $f(\xi)$, where $f(\infty) = f_0$ and $f(\xi_L) = f(\xi_H) = f_1$.

Trade policy. All firms in each good g face the same tariff, τ_g, which can take one of two values: the NNTR (Column Two) tariff, τ_{g2}, or the MFN (Column One) tariff, τ_{g1}. The tariff regime is an aggregate state variable that follows a first-order, time-varying Markov process. If the current tariff regime is NNTR, the probability that the regime in the next period will be MFN is $\omega_{21,t}$. Conversely, the probability of switching from MFN tariffs to NNTR tariffs is $\omega_{12,t}$. The fact that the trade-policy regime is an aggregate state means all goods (and all firms in each good) face the same regime and the same transition probabilities at each point in time. In the benchmark model, firms know the entire path of $\{\omega_{21,t}\}_{t=0}^{\infty}$ and $\{\omega_{12,t}\}_{t=0}^{\infty}$. We consider alternative information structures in section 4.4.

Firm optimization. The firm's export status is determined in the prior period. The firm maximizes current-period profits by choosing its price, taking as given its residual demand

and the wage, w,

$$\pi_{gt}(z_t, \xi_t, \tau_{gt}) = \max_{p,\,\ell} p\, d_{gt}(p, \xi_t, \tau_{gt}) - w_t \ell \tag{9}$$

$$\text{s.t. } \; z_t \ell \geq d_{gt}(p, \xi_t, \tau_{gt}) \xi_t. \tag{10}$$

The value of a firm that chooses to export at $t+1$ is

$$V^1_{gt}\left(z_t, \xi_t, \tau_{gt}\right) = -f(\xi_t) + \frac{\delta(z_t)}{1+r} \mathop{\mathbb{E}_{\mathrm{t}}}_{z,\xi,\tau_g} V_{g,t+1}\left(z_{t+1}, \xi_{t+1}, \tau_{g,t+1}\right), \tag{11}$$

where r is the interest rate used to discount future profit. The value of a firm that chooses not to export at $t+1$ is

$$V^0_{gt}\left(z_t, \xi_t, \tau_{gt}\right) = \frac{\delta(z_t)}{1+r} \mathop{\mathbb{E}_{\mathrm{t}}}_{z,\xi,\tau_g} V_{t+1}\left(z_{t+1}, \xi_{t+1}, \tau_{g,t+1}\right), \tag{12}$$

and the value of the firm is

$$V_{gt}\left(z_t, \xi_t, \tau_{gt}\right) = \pi_{gt}(z_t, \xi_t, \tau_{gt}) + \max\left\{V^1_{gt}\left(z_t, \xi_t, \tau_{gt}\right), V^0_{gt}\left(z_t, \xi_t, \tau_{gt}\right)\right\}. \tag{13}$$

The break-even exporter $\bar{z}_{gt}(\xi)$ is indifferent between exporting and not exporting:

$$V^1_{gt}\left(\bar{z}_{gt}(\xi), \xi, \tau_{gt}\right) = V^0_{gt}\left(\bar{z}_{gt}(\xi), \xi, \tau_{gt}\right). \tag{14}$$

Aggregation. The solutions to these programming problems yield decision rules over export participation and, conditional on exporting, export volumes. The decision rules determine the laws of motion over firms. The measure of firms that do not export is $\varphi^0_{gt}(z, \xi)$, and the measure of exporters is $\varphi^1_{gt}(z, \xi)$. We use these measures to compute aggregate trade volumes,

$$EX_{gt} = \sum_{\xi \in \{\xi_L, \xi_H\}} \int_z p\left(z, \xi, \tau_{gt}\right) y\left(z, \xi, \tau_{gt}\right) \varphi^1_{gt}\left(z, \xi\right) dz. \tag{15}$$

4.2 Calibration

The model is calibrated to reproduce facts about the historical dynamics of U.S. imports from China. We initialize the model in 1970 with zero exporters (that is, all firms have $\xi = \infty$) in all goods to capture the transition from autarky, and then feed in the realized path of tariffs (NNTR rates during 1971–1979 and MFN rates from 1980 onward). We then choose the model's parameters so that its transition dynamics match the facts documented in section 3 and other salient facts about export participation dynamics reported elsewhere in the literature. The calibrated parameter values are summarized in Table 7.

Assigned parameters. Several of the model's parameters are assigned externally. A model period is one year. The wage is normalized to one and the interest rate used to discount future profits is four percent. We take the time series for MFN and NNTR tariffs, τ_{g1} and τ_{g2}, directly from the data. The parameters that govern the firm's productivity and survival processes are taken from Alessandria et al. (2021a). To eliminate the role of the elasticity of substitution in the size distribution of firms, we assume producer productivity is $z = \frac{1}{\theta-1}\log a$. A firm's productivity follows

$$\log a_{t+1} = \rho_z \ln a_t + \varepsilon, \qquad \varepsilon \overset{iid}{\sim} N(0, \sigma_z^2), \tag{16}$$

which we convert to a Markov-chain approximation for computation. Firms are subject to exogenous death shocks that depend on the firm's productivity; the probability of death is $1 - \delta(a) = \max\left\{0, \min\left\{e^{-\delta_0 a} + \delta_1, 1\right\}\right\}$.

Parameters calibrated to facts about exporter dynamics. The parameters that govern fixed and variable exporting costs are calibrated so that the model's terminal steady state—in which tariffs have remained at MFN rates for many years—matches statistics on export participation dynamics documented by Alessandria et al. (2021a). The sunk entry cost, f_0, is chosen so that 22.3 percent of firms export. The continuation cost, f_1, is chosen so that 17.0 percent of exporters exit. The low variable exporting cost, ξ_L, is normalized to one and the high variable exporting cost, ξ_H, is chosen so that the average export entrant is half as large as the average incumbent exporter (as measured by sales).

Parameters calibrated to short- and long-run trade elasticities. The elasticity of demand, θ, and the probability of retaining the current variable export cost, ρ_ξ, are chosen so that estimating the unrestricted error-correction model (3) on simulated data from the model reproduces the short-run and long-run trade elasticities for China of 2.3 and 8.06, respectively, as reported in Table 2. We find $\theta = 3.55$ and $\rho_\xi = 0.87$.[24]

Parameters calibrated to annual NNTR-gap elasticities. The key parameters in our model are the probabilities of switching between tariff regimes, $\omega_{12,t}$ and $\omega_{21,t}$. We calibrate these parameters so that estimating equation (5) on simulated data from our model reproduces the baseline annual NNTR-gap elasticities reported in column 1 of Table 5. The probability that the economy moves from the NNTR rates to the MFN rates, $\omega_{21,t}$, is assumed to be constant and is set to match the average NNTR gap during the 1974–1979 period. We allow the probability of reverting from the MFN tariff rates back to the NNTR rates, $\omega_{12,t}$, to vary over time, and choose it to match the annual elasticity coefficients from 1980 onward.[25] These parameters are identified by the way export participation in high-gap industries responds to trade-policy uncertainty relative to low-gap industries. During the 1974–1980 period, a higher probability of moving from NNTR tariff rates to MFN rates boosts export participation in high-gap industries more than in low-gap industries, which increases the elasticity coefficient during this period.[26] After 1980, a higher probability of moving from MFN tariff rates back to NNTR rates reduces export participation in high-gap industries more than in low-gap industries, so the model interprets the decline in the magnitude of the elasticity coefficient during this period as a decline in the probability of moving back to NNTR status.

[24]Note our calibrated demand elasticity is greater than the estimated short-run elasticity. We spread the intensive-margin effect of the 1980 liberalization over two years (1980–1981) by assuming only half of each industry's exporters get access to MFN tariff rates in 1980. We use this approach to capture the fact that the NNTR-gap elasticity rises in a linear fashion from 1979 through 1981, as shown in Figure 6.

[25]We HP-filter the coefficients from 1981 onward to smooth out temporary spikes (such as 1984), but we use the 1980 coefficient without modification because it plays a dominant role in identifying the demand elasticity, θ, which governs the short-run elasticity. The calibrated model matches this smoothed series almost perfectly.

[26]We have also experimented with a version of the model that includes a probability of moving back to autarky. A higher probability of moving to autarky, however, has the same effect on the NNTR gap as a higher probability of moving to MFN status. Moving to autarky hurts export participation more in low-gap industries than in high-gap industries, because the former experienced larger initial liberalizations than the latter.

In panel (a) of Figure 11, we plot the annual NNTR-gap elasticities from the model and the data. The elasticities from the data are from Figure 6. Overall, the model fits the data quite well during all three periods of U.S.-China trade history: the 1970s, before China was granted NTR status; 1980–2000, after this status was unilaterally granted; and 2001 onward, when China was a WTO member.

4.3 Estimates of Expectations of U.S.-China Trade Policy

Panel (b) of Figure 11 plots the main results of our calibration exercise: the estimated probabilities of switching between policy regimes: We assume the probability of transitioning from the NNTR to the MFN regime is constant and estimate it to be about 0.25. This assumption allows the model to naturally match the slight downward drift in the elasticity of trade to the NNTR gap from 1974 to 1979. The probability that the economy switches back to the NNTR regime is initially 0.72 in 1980, rises to 0.80 in 1981, and then begins to fall throughout the 1980s and early 1990s. By 1993, the regime-switching probability is less than eight percent. A temporary increase occurs during 1994–1996, after which the probability continues to fall throughout the late 1990s and 2000s. By 2008, the end of our observation period, the probability of moving back to the NNTR regime falls to 3.4 percent. Note no discrete drop or accelerated decline occurs in this probability when China gains PNTR status in 2001 (in fact, the switching probability rises slightly, from 5.3 percent to 6.3 percent, during 2000–2002), which reflects the observation in section 3.3.2 that this change in status had little impact on the annual elasticity of trade to the NNTR gap. This finding indicates gaining PNTR status did not materially alter Chinese exporters' beliefs about future U.S. trade policy; switching back to the NNTR regime was viewed as extremely unlikely even before China joined the WTO.

Another way to interpret these results is to ask: what is the expected present value of tariffs that Chinese exporters face at each point in time? Panel (c) of Figure 11 plots the mean expected present value of tariffs across goods in the model,

$$\tau_t^{PV} = \frac{1}{G}\left\{\sum_{g=1}^{G}(1-\beta)\left(\sum_{s=t}^{\infty}\beta^{s-t}\mathbb{E}_t[\tau_{gs}]\right)\right\}, \tag{17}$$

alongside the mean applied tariff (NNTR before 1979 and MFN from 1980 onward). Whereas the realized path of applied tariffs falls sharply in 1980 when China gains NTR status, and then falls slightly throughout the 1980s and 1990s due to continued reforms to U.S. MFN tariff rates, the expected present value of tariffs falls gradually but steadily throughout the entire period. We find only a small discrete drop in 1980, and the expected present value remains well above the applied MFN rate even after China joins the WTO in 2001.

The behavior of the expected tariff helps us understand the gradual adjustment of export volumes to the abrupt decrease in the current tariff. The large expected tariff suppressed entry of Chinese firms in the export market. As the expected tariff fell, entry increased and aggregate trade volumes grew. The uncertainty over future tariffs mattered the most in the 1980s and 1990s.

4.4 Alternative Expectations

In our benchmark model, firms do not know which trade-policy regimes they will face in the future, but they know how the probabilities of switching between regimes will evolve over time. Here, we compare our benchmark model with two alternatives: a counterfactual in which no trade policy uncertainty is present and another in which uncertainty exists, but firms do not anticipate changes in regime-switching probabilities. We also compare our model-implied probabilities of reverting from MFN to NNTR tariffs with those obtained via Bayesian learning.

4.4.1 No Trade-Policy Uncertainty

In the counterfactual without trade-policy uncertainty, the path of tariffs is deterministic. Starting in 1970, firms know they will face NNTR tariffs during 1971–1979 and MFN tariffs from 1980 onwards. Because the realized path of tariffs in the benchmark model with TPU is the same as in the no-TPU counterfactual, differences in outcomes between the two models are due solely to the presence of uncertainty in the former.

Panel (d) of Figure 11 plots aggregate trade flows in the benchmark model and the counterfactual without trade policy uncertainty. In the benchmark, Chinese exports to the United States grow modestly after the embargo is lifted in 1971, jump slightly when the NTR liberalization occurs in 1980, and continue to grow gradually throughout the 1980s, 1990s, and 2000s. The jump in exports between 1979 and 1981 is the growth in the exports of firms that were already exporting when the liberalization took effect. The gradual growth that follows comes from nonexporting firms choosing to enter the export market and firms that are already exporting becoming better exporters (firms moving from ξ_H to ξ_L). Although the average tariff rate declines further as MFN tariffs are changed, much of the protracted growth we observe is the slow adjustment to the liberalization in 1980.

In the counterfactual without trade-policy uncertainty, exports grow much faster immediately after the embargo is lifted, and grow more slowly after 1985. The rapid growth during the 1970s is driven by firms that choose to start exporting in anticipation of the 1980 liberalization. The slower growth in the latter part of the observation period indicates the counterfactual has largely converged to a steady state by the late 1980s. The vertical distance between the two lines represents a measure of the effect of trade-policy uncertainty on exports in a given period. This effect is largest immediately after the 1980 liberalization, consistent with our empirical findings about the evolution of the NNTR gap and our model's estimate of how the probability of reverting back to NNTR status fell over time. We see little change in this effect when China joins the WTO in 2001. In fact, trade-policy uncertainty still depresses exports several years afterward. This finding is explained by panel (c) of Figure 11, which shows the expected present value of tariffs in our benchmark model is still noticeably higher at the end of our sample than in the counterfactual without trade policy uncertainty.

4.4.2 Unanticipated Changes in Trade-Policy Uncertainty

In our second alternative, we assume the trade-policy regime follows a Markov process as in the benchmark, but that in each period, firms believe the current transition probabilities will remain in force forever—firms are surprised each period when these probabilities change. In this version of the model, we recalibrate the transition probabilities to match the annual NNTR-gap coefficients while leaving all other aspects of the calibration unchanged. As in the first alternative, the realized path of tariffs is the same as in the benchmark, so any differences in outcomes are due to differences in firms' expectations.

Panel (b) of Figure 11 shows the calibrated transition probabilities in this version of the model (labeled "surprises") are very similar to the benchmark probabilities. The initial probability of losing MFN status in 1980 is slightly higher but falls slightly quicker thereafter. Panel (c) of Figure 11 shows these small differences in probabilities imply the expected present value of tariffs is similar throughout the sample period in this version of the model to the benchmark, and Panel (d) shows they imply similar paths of aggregate trade.

These two setups differ in their assumptions about firms' expectations over the probabilities of switching between policy regimes. These different expectations yield quite similar paths of probabilities in large part because they yield similar present values of tariffs (panel (c) of Figure 11) in each period. This finding suggests our approach provides tight bounds on these probabilities and on their economic effects.

4.4.3 Bayesian Learning

In our model, firms do not update their beliefs about the probabilities of switching between regimes; they simply take at face value the probabilities that are "announced" by the modeler. Here, we ask how the probabilities that we obtain from our calibration exercise compare with the posterior beliefs that a Bayesian agent would form after observing the economy remains in the MFN regime year after year from 1980 onward.

We focus this analysis on the probability of losing MFN status after the 1980 liberalization, ω_{12}. We assume Bayesian agents have beta-distributed prior beliefs about this probability when the liberalization occurs in 1980:

$$p^{prior}(\omega_{12}|a,b) = \frac{\Gamma(a+b)}{\Gamma(a)+\Gamma(b)}\omega_{12}^{a-1}(1-\omega_{12})^{b-1}. \tag{18}$$

The parameters a and b of this distribution control the mean and the degree of confidence in this value. For example, $a = b = 1$ is the uniform distribution that has a mean of 0.5 but places equal weight on all possible values of ω_{12}, whereas the beta distribution with $a = b = 5$ has the same mean but is tightly concentrated around that value. This conjugate prior distribution is convenient because the mean posterior after observing n successive periods in which MFN status is retained is given by the simple expression $a/(a+b+n)$.

We consider a range of priors that all have the same mean as the initial 1980 probability in the model but with more or less dispersion around this value, that is, priors that are more or less confident that this initial probability is correct. This setup allows us to determine whether agents in our model "learn" faster or slower than a Bayesian agent would. Specifically, for each $b = 1, 2, \ldots, 5$, we set a so that the mean prior, given by $a/(a + b)$, equals 0.72. The prior with $b = 1$ represents an agent with little confidence in this value, whereas the prior with $b = 5$ represents a highly confident agent. Panel (a) of Figure 12 plots the density functions of each of the prior beliefs that we consider.

Panel (b) of Figure 12 plots the model-implied probabilities of losing MFN status against the evolution of the mean posteriors associated with each of these priors as agents observe more and more successive periods in which MFN status is retained. During 1980–1985, the Bayesian posteriors fall faster than our model-implied probabilities, which is consistent with the delay in growth in the NNTR-gap coefficient during the early 1980s documented in section 3.3.2. After 1985, however, this pattern is reversed, and by the late 1990s the model-implied probability of losing MFN status is lower than the mean posterior for all of the priors we consider— even very widely dispersed priors that quickly incorporate new information (such as $b = 1$).

4.5 Recovering the Long-Run Trade Elasticity

In addition to measuring the extent of trade-policy uncertainty, we can use our model to recover the long-run effect of trade policy on trade flows in the absence of uncertainty—the long-run trade elasticity—which is the key moment in most empirical and theoretical trade studies. We find the long-run trade elasticity in our model without uncertainty is higher than what we estimate in the data, using either cross-sectional or time-series methods, and that it is increasing in the initial tariff. The large differences between our empirical estimates of the long-run trade elasticity and the model's long-run trade elasticity point to the challenges of measuring this object in the data in the face of uncertainty about future trade policy.

To solve for our model's long-run trade elasticity, we eliminate all tariff uncertainty and solve for the stationary equilibrium associated with the full range of tariffs observed in the historical data, from NNTR rates to modern MFN rates. We measure it as the change in the steady-state level of trade in a given good caused by a marginal change in the tariff on that good, as a function of the initial tariff. We plot the results of this exercise in Figure 13.

When tariffs are close to zero, the trade elasticity is close to 15, and it rises to about 18.5 at the highest tariff levels.

Our model's long-run trade elasticity of 15–18 is larger than those recovered from our empirical approaches. As reported in Table 2, the elasticity estimated in the pooled cross section of U.S. imports from China is 6–7. This estimate, however, is obtained from the entire historical sample of U.S.-China trade, which includes both short-run and long-run responses to trade reforms, and therefore understates the true long-run elasticity. Using panel methods, we find a slightly higher trade elasticity of 8 to 8.5 using the error-correction model and an elasticity of about 10 when using the NNTR gap in the 1970s. Our model, which is calibrated to reproduce these estimates, indicates that even these dynamic methods understate the true long-run response of trade to permanent, deterministic changes in trade policy.

5 Conclusions

We study, empirically and theoretically, China's integration into the U.S. market after the long-standing embargo on Chinese goods was lifted in 1971. We find the dynamics of integration are consistent with substantial uncertainty about the future path of tariffs in the late 1970s and early 1980s. Compared with the earlier period, our results imply much less policy uncertainty in the late 1990s in the lead-up to China being granted permanent normal trade relations in 2000. The substantial drop in the probability of NTR reversal in the mid 1980s, combined with the gradual trade expansion common to trade liberalizations, suggests much of the growth in trade in high NNTR-gap products in the 1990s was due to the prior liberalization and not the future outlook on trade policy.

Our approach to estimating a forward-looking path of trade policy leverages unique aspects of U.S. policy toward China in which the change in trade policy is known and heterogeneous across products, but the likelihood and timing is unknown but common across products. Our analysis could be extended to consider more contemporary events such as Brexit, the U.S.-China trade war, safeguards, and domestic content requirements, as well as traditional protectionist measures such as antidumping duties. In all these cases, the size and timing of the reforms are uncertain, but by interpreting trade flows through a dynamic model, we can discipline the process for these possible trade-policy outcomes.

Our analysis produces estimates of key time-varying trade elasticities and probabilities of policy reversal that should be useful in disciplining general-equilibrium models of trade policy and trade dynamics. Reconsidering the aggregate effects of China's global integration, taking into account the dynamics of trade policy we have identified, would be interesting. An important challenge to a general-equilibrium analysis is accounting for non-tariff barriers such as quotas or safeguards, because these key aspects of China trade policy were changed by the United States in the late 1990s and the 2000s.

Finally, that trade depends on past, current, and future changes in trade policy suggests we need to rethink our notions of changes in trade policy, the trade elasticity, and the measurement of these trade elasticities. Alessandria et al. (2021b) build on our findings here to show how to measure the trade response to unanticipated and anticipated changes in trade policy.

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Table 1: Summary Statistics (percent)

SITC 1-Digit	NNTR Rate		U.S. Export Share		Applied Duties			
					1979		2001	
	Mean	Std.	1979	2001	China	NTR	China	NTR
0 Food and live animals (food)	17	15	10	1	12	5	6	2
1 Beverages and tobacco	48	41	0	0	31	11	4	3
2 Crude materials, inedible, except fuels	11	15	13	1	6	1	1	0
3 Mineral fuels, lubricants and related	3	8	15	0	2	0	2	0
4 Animal and vegetable oils, fats and waxes	14	9	1	0	4	2	2	2
5 Chemical and related products	31	19	10	2	16	4	2	2
6 Manufactured goods	41	21	13	12	33	8	4	4
7 Machinery and transport equip.	33	10	0	36	29	5	2	1
8 Misc. manufactured articles	50	25	37	49	38	14	5	4
9 Commodities and transactions, n.e.c.	29	19	0	0	29	0	1	1
Average	28	18			20	5	3	2

Notes: NNTR Rate is the median HS-8 NNTR rate from Feenstra et al. (2002) at the 5-digit SITC level. *U.S. Export Share* is the share of total Chinese exports shipped to the United States. *Applied Duties* is the mean 5-digit SITC tariff (duties over FOB value). NTR refers to the set of countries that never received NNTR treatment by the United States nor had a FTA with it (see footnote 14). When computing the mean applied duties of NTR countries, only SITC codes with nonzero U.S. imports from China are considered.

Table 2: Short- and Long-Run Trade Elasticity: Baseline

	Cross-Section		ECM Rest.	ECM Unrest.
	ν_{jgt}	ν_{jgt}	ν_{jgt}	$\Delta\nu_{jgt}$
$\mathbb{1}\{j \neq China\}\tau_{jst}$	-2.66^{***}	-2.60^{***}	-1.86^{***}	
	(0.22)	(0.26)	(0.19)	
$\mathbb{1}\{j = China\}\tau_{jgt}$	-6.69^{***}	-6.80^{***}	-2.84^{***}	
	(0.32)	(0.38)	(0.24)	
$\mathbb{1}\{j \neq China\}v_{jg,t-1}$			0.53^{***}	-0.47^{***}
			(0.00)	(0.00)
$\mathbb{1}\{j = China\}v_{jg,t-1}$			0.63^{***}	-0.37^{***}
			(0.01)	(0.01)
$\mathbb{1}\{j \neq China\}\Delta\tau_{jgt}$				-2.02^{***}
				(0.24)
$\mathbb{1}\{j = China\}\Delta\tau_{jgt}$				-2.29^{***}
				(0.41)
$\mathbb{1}\{j \neq China\}\tau_{jg,t-1}$				-1.61^{***}
				(0.16)
$\mathbb{1}\{j = China\}\tau_{jg,t-1}$				-2.95^{***}
				(0.22)
Long-run NTR Countries			-3.97^{***}	-3.45^{***}
			(0.41)	(0.35)
Long-run China			-7.77^{***}	-8.06^{***}
			(0.65)	(0.57)
Long-/Short-Run NTR Countries			2.14	1.71
Long-/Short-Run China			2.74	3.51
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	1,185,475	934,554	934,554	934,554
Adj. R-squared	0.78	0.79	0.86	0.27

Notes: Results in columns 1 and 2 are estimated as the log of imports on the log of tariffs. Column 2 restricts the sample to be the same as in columns 3 and 4 (requiring non-missing lagged imports). Results in column 3 are from estimating (2), and those in column 4 are from estimating (3). In column 2, the coefficients on τ_{jgt} capture the short-run elasticity. In column 4, the coefficients on $\Delta\tau_{jgt}$ capture the short-run elasticity. The long-run elasticity in column 2 is the coefficient on τ_{jgt} divided by one minus the coefficient on $v_{jg,t-1}$. The long-run elasticity in column 3 is the coefficient on $\tau_{jg,t-1}$ divided by the coefficient on $v_{jg,t-1}$. Standard errors in parentheses are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 3: Short- and Long-Run Trade Elasticity: Robustness

	Baseline	Shipping	Life cycle	All	Full	Balanced	NAFTA	NAFTA+MFA
$\mathbb{1}\{j \neq China\}v_{jg,t-1}$	-2.02^{***}	-1.90^{***}	-2.02^{***}	-2.08^{***}	-2.18^{***}	-2.08^{***}	-1.93^{***}	-2.73^{***}
	(0.13)	(0.13)	(0.13)	(0.12)	(0.10)	(0.15)	(0.54)	(0.53)
$\mathbb{1}\{j = China\}v_{jg,t-1}$	-2.29^{***}	-2.23^{***}	-2.29^{***}	-2.31^{***}	-2.66^{***}	-2.35^{***}	-2.21^{***}	-2.69^{***}
	(0.38)	(0.39)	(0.37)	(0.38)	(0.38)	(0.41)	(0.40)	(0.41)
$\mathbb{1}\{j \neq China\}\Delta\tau_{jgt}$	-0.47^{***}	-0.47^{***}	-0.47^{***}	-0.46^{***}	-0.46^{***}	-0.44^{***}	-0.34^{***}	-0.34^{***}
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)
$\mathbb{1}\{j = China\}\Delta\tau_{jgt}$	-0.37^{***}	-0.37^{***}	-0.37^{***}	-0.37^{***}	-0.37^{***}	-0.31^{***}	-0.35^{***}	-0.35^{***}
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
$\mathbb{1}\{j \neq China\}\tau_{jg,t-1}$	-1.61^{***}	-1.49^{***}	-1.62^{***}	-1.72^{***}	-1.89^{***}	-1.56^{***}	-2.00^{***}	-2.63^{***}
	(0.11)	(0.11)	(0.11)	(0.09)	(0.08)	(0.11)	(0.37)	(0.32)
$\mathbb{1}\{j = China\}\tau_{jg,t-1}$	-2.95^{***}	-2.96^{***}	-2.96^{***}	-2.96^{***}	-2.85^{***}	-2.66^{***}	-2.56^{***}	-2.55^{***}
	(0.30)	(0.30)	(0.30)	(0.29)	(0.26)	(0.27)	(0.30)	(0.28)
Shipping Costs_{jgt}		-2.78^{***}						
		(0.05)						
Long-run China	-8.06^{***}	-8.01^{***}	-8.05^{***}	-8.06^{***}	-7.67^{***}	-8.51^{***}	-7.34^{***}	-7.31^{***}
	(0.79)	(0.79)	(0.78)	(0.77)	(0.67)	(0.82)	(0.85)	(0.77)
Long-Run Other	-3.45^{***}	-3.15^{***}	-3.45^{***}	-3.75^{***}	-4.14^{***}	-3.55^{***}	-5.88^{***}	-7.69^{***}
	(0.23)	(0.23)	(0.23)	(0.20)	(0.18)	(0.26)	(1.08)	(0.94)
Long-/Short-Run China	3.51	3.60	3.52	3.49	2.89	3.61	3.32	2.71
Long-/Short-Run Other	1.71	1.65	1.71	1.80	1.89	1.71	3.04	2.82
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	934,554	934,554	934,554	1,053,446	1,232,395	569,854	92,614	105,738
Adj. R-squared	0.27	0.28	0.27	0.26	0.26	0.26	0.24	0.25

Notes: All results are obtained estimating (3). The short-run elasticity is captured by the coefficient on $\Delta\tau_{jgt}$. The long-run elasticity is the coefficient on $\tau_{jg,t-1}$ divided by the coefficient on $v_{jg,t-1}$. The *Life cycle* model includes life-cycle controls for the first, second, penultimate, and last year that a good is traded, as well as the good's age and age squared. The *All* model includes all countries, the *Full* model further includes goods affected by the MFA quotas, and the *Balanced* model is restricted to goods with non-zero U.S.-China trade before 1981. Standard errors in parentheses are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 4: Elasticity of Trade to NNTR Gap

	ν_{jgt}					
	PS (2016)	1974–08	NTR	Ex. Tariffs	NNTR	App. NNTR
$\mathbb{1}_{\substack{t>2000\\ j=China}} NNTRGap_g$	-0.92^{***}	-2.50^{***}	-3.36^{***}	-3.60^{***}		
	(0.22)	(0.28)	(0.32)	(0.33)		
$\mathbb{1}_{\substack{t>2000\\ j=China}} NNTR_g$					-3.09^{***}	
					(0.32)	
$\mathbb{1}_{\substack{t>2000\\ j=China}} AppNNTR_g$						-4.13^{***}
						(0.42)
τ_{jgt}	-3.28^{***}	-3.20^{***}	-2.82^{***}			
	(0.14)	(0.11)	(0.14)			
Period	'92–'07	'74–'08	'74–'08	'74–'08	'74–'08	'74–'08
Countries	All	All	NTR	NTR	NTR	NTR
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	781,960	1,453,687	1,090,460	1,090,460	1,090,460	614,557
Adj. R-squared	0.84	0.78	0.78	0.78	0.78	0.80

Notes: All estimates are obtained from variations of (4). Column 1 estimates the effect of the NNTR Gap on Chinese imports after its PNTR access using the same period as in Pierce and Schott (2016) but at the SITC aggregation level (see equation 5 in Pierce and Schott (2016)). Column 2 estimates the same equation using the sample period 1974 to 2008. Column 3 further includes only NTR countries and China, and column 5 excludes tariffs from (4). Columns 5 and 6 use the NNTR rate in 2001 and the applied NNTR rate instead of the NNTR gap. Standard errors in parentheses are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 5: Annual Elasticity to NNTR Gap: Robustness

	ν_{jgt}				
	Baseline	Anticipation	Shipping	Tariffs	Life cycle
$\mathbb{1}_{t=t'}\mathbb{1}_{j=China}X_g$					
1974	-10.2^{***}	-9.83^{***}	-10.2^{***}	-8.41^{***}	-10.2^{***}
1975	-9.69^{***}	-10.2^{***}	-9.73^{***}	-7.95^{***}	-9.75^{***}
1976	-10.7^{***}	-10.4^{***}	-10.7^{***}	-8.72^{***}	-10.8^{***}
1977	-10.2^{***}	-10.5^{***}	-10.3^{***}	-8.31^{***}	-10.4^{***}
1978	-10.4^{***}	-9.88^{***}	-10.5^{***}	-8.38^{***}	-10.4^{***}
1979	-10.9^{***}	-9.84^{***}	-11.0^{***}	-9.04^{***}	-11.0^{***}
1980	-9.27^{***}	-9.08^{***}	-9.34^{***}	-9.07^{***}	-9.35^{***}
1981	-7.50^{***}	-7.42^{***}	-7.53^{***}	-7.36^{***}	-7.57^{***}
1982	-7.69^{***}	-7.67^{***}	-7.71^{***}	-7.56^{***}	-7.81^{***}
1983	-7.84^{***}	-7.46^{***}	-7.83^{***}	-7.74^{***}	-7.89^{***}
1984	-6.44^{***}	-6.76^{***}	-6.44^{***}	-6.36^{***}	-6.52^{***}
1985	-7.30^{***}	-7.08^{***}	-7.29^{***}	-7.21^{***}	-7.41^{***}
1986	-6.97^{***}	-6.74^{***}	-6.98^{***}	-6.92^{***}	-7.04^{***}
1987	-6.16^{***}	-5.98^{***}	-6.22^{***}	-6.17^{***}	-6.28^{***}
1988	-4.80^{***}	-4.58^{***}	-4.84^{***}	-4.81^{***}	-4.89^{***}
1989	-4.20^{***}	-3.93^{***}	-4.08^{***}	-4.04^{***}	-4.29^{***}
1990	-3.42^{***}	-3.45^{***}	-3.39^{***}	-3.36^{***}	-3.51^{***}
1991	-2.69^{***}	-2.48^{***}	-2.76^{***}	-2.71^{***}	-2.80^{***}
1992	-2.20^{***}	-2.12^{***}	-2.24^{***}	-2.20^{***}	-2.29^{***}
1993	-1.70^{***}	-1.63^{***}	-1.76^{***}	-1.70^{***}	-1.79^{***}
1994	-1.81^{***}	-1.65^{***}	-1.75^{***}	-1.69^{***}	-1.89^{***}
1995	-1.79^{***}	-1.74^{***}	-1.79^{***}	-1.75^{***}	-1.89^{***}
1996	-1.62^{***}	-1.56^{***}	-1.62^{***}	-1.54^{***}	-1.72^{***}
1997	-1.79^{***}	-1.58^{***}	-1.69^{***}	-1.68^{***}	-1.88^{***}
1998	-1.49^{***}	-1.52^{***}	-1.40^{***}	-1.39^{***}	-1.56^{***}
1999	-1.15^{***}	-1.12^{***}	-1.10^{**}	-1.09^{**}	-1.21^{***}
2000	-0.68^{*}	-0.34	-0.71^{*}	-0.71^{*}	-0.73^{*}
2001	-0.24	-0.31	-0.31	-0.31	-0.29
2002	-0.61^{*}	-0.59^{*}	-0.61^{*}	-0.61^{*}	-0.67^{*}
2003	-0.93^{***}	-0.68^{**}	-0.86^{**}	-0.87^{**}	-0.96^{***}
2004	-0.27	-0.37	-0.26	-0.26	-0.31
2005	-0.59^{*}	-0.51^{*}	-0.57^{*}	-0.57^{*}	-0.62^{*}
2006	-0.25	-0.30	-0.24	-0.24	-0.26
2007	-0.11	0	-0.16	-0.16	-0.13
$\Delta\tau_{jg,t+1}$		1.05^{***}			
τ_{jgt}				-2.51^{***}	
Shipping Cost_{jgt}			-3.03^{***}	-3.02^{***}	
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	1,090,460	862,714	1,090,460	1,090,460	1,090,460
Adj. R-squared	0.78	0.78	0.78	0.78	0.78

Notes: All estimates are obtained using variations of (5). Column 2 — *Anticipation* — includes the lead change in tariffs to control for some of the anticipation to the 1980 NTR liberalization. Column 3 — *Shipping* — includes shipping costs. Column 4 — *Anticipation* — further includes applied duties. Column 5 — *Life cycle* — includes life-cycle controls for the first, second, penultimate, and last year that a good is traded, as well as the good's age and age squared. Standard errors in parentheses are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 6: Annual Elasticity to NNTR Gap: Robustness (continued)

	ν_{jgt}				
	All	Full	Balanced	NNTR	Applied
$\mathbb{1}_{t=t'}\mathbb{1}_{j=China}X_g$					
1974	−9.88***	−8.88***	−10.1***	−8.23***	−9.43***
1975	−9.54***	−8.37***	−9.60***	−8.37***	−9.34***
1976	−10.5***	−8.16***	−10.6***	−9.31***	−9.42***
1977	−10.1***	−8.00***	−10.2***	−9.05***	−8.92***
1978	−10.3***	−7.50***	−10.3***	−9.01***	−9.01***
1979	−10.8***	−7.64***	−10.9***	−9.49***	−9.25***
1980	−9.13***	−6.17***	−9.19***	−8.27***	−7.77***
1981	−7.38***	−4.83***	−7.57***	−6.30***	−6.69***
1982	−7.58***	−4.82***	−7.45***	−6.77***	−6.82***
1983	−7.67***	−4.48***	−7.39***	−6.67***	−6.38***
1984	−6.33***	−3.95***	−6.48***	−5.54***	−5.95***
1985	−7.18***	−4.63***	−6.83***	−6.56***	−5.96***
1986	−6.87***	−4.15***	−6.88***	−6.17***	−5.53***
1987	−6.09***	−3.82***	−6.10***	−5.53***	−5.13***
1988	−4.76***	−3.20***	−4.51***	−4.41***	−3.76***
1989	−4.08***	−2.54***	−3.81***	−3.66***	−2.76***
1990	−3.27***	−2.29***	−3.07***	−3.09***	−2.56***
1991	−2.60***	−1.73***	−2.55***	−2.27***	−2.00***
1992	−2.13***	−1.41***	−2.58***	−1.83***	−1.63***
1993	−1.63***	−0.90**	−2.07***	−1.31***	−1.58***
1994	−1.76***	−1.19***	−2.14***	−1.41***	−1.54***
1995	−1.67***	−1.18***	−1.90***	−1.52***	−1.10**
1996	−1.55***	−1.18***	−1.53***	−1.37***	−1.07**
1997	−1.78***	−1.17***	−1.85***	−1.50***	−1.17**
1998	−1.51***	−1.25***	−1.01**	−1.40***	−0.90**
1999	−1.19***	−0.88**	−1.29***	−0.94**	−1.10**
2000	−0.68*	−0.55	−0.91**	−0.56	−0.68*
2001	−0.27	−0.21	−0.35	−0.24	−0.25
2002	−0.63*	−0.29	−0.76**	−0.51	−0.34
2003	−0.99***	−0.52*	−0.83**	−0.85***	−0.56
2004	−0.34	0.013	−0.30	−0.20	0.15
2005	−0.64*	0.058	−0.29	−0.65*	0.32
2006	−0.37	0.31	−0.57*	−0.21	−0.33
2007	−0.14	−0.17	0.13	−0.15	0.28
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	1,223,264	1,453,687	658,635	1,090,460	614,557
Adj. R-squared	0.78	0.78	0.80	0.78	0.80

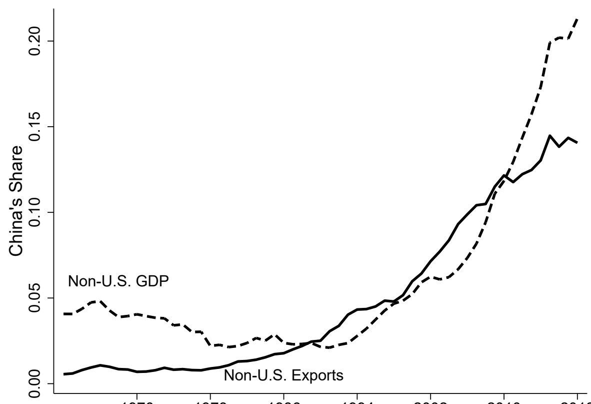
Notes: All estimates are obtained using variations of (5). X_g is the good-level NNTR Gap in 2001, except in column 4, in which it is the NNTR rate, and in column 5, in which it is the applied NNTR rate, calculated as the good-level tariff rate applied to Chinese imports between 1974 and 1979. Countries included in the baseline sample are China and countries with NTR, excluding Mexico and Canada. Column *All* includes all countries. Column *Balanced* includes only the goods in which U.S.-China trade was non-zero before 1981. Standard errors in parentheses are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table 7: Calibrated Parameter Values

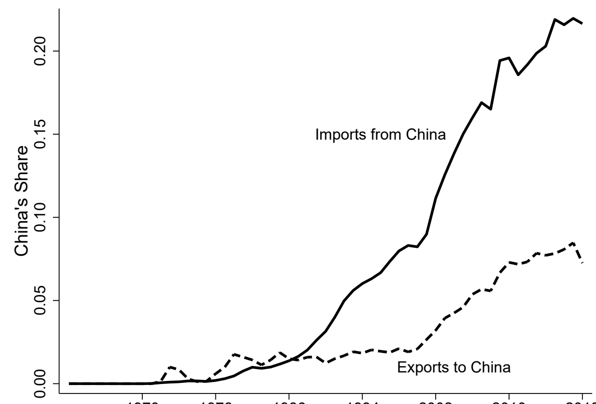
Parameter	Meaning	Value	Source/target
(a) Assigned			
w	Wage	1	Normalization
r	Interest rate	4 pct.	Standard
σ_z	Variance of productivity	1.32	Alessandria et al. (2021a)
ρ_z	Persistence of productivity	0.65	Alessandria et al. (2021a)
δ_0	Correlation of survival with productivity	21.04	Alessandria et al. (2021a)
δ_1	Minimum death probability	0.023	Alessandria et al. (2021a)
τ_{g1}	NNTR tariff	Varies	Data
τ_{g2}	MFN tariff	Varies	Data
(b) Calibrated to exporter life cycle			
f_0	Entry cost	0.50	Export participation rate = 22 pct.
f_1	Continuation cost	0.30	Exit rate = 17 pct.
ξ_H/ξ_L	High iceberg cost	2.60	Avg. entrant sales/avg. incumbent sales = 0.5
(c) Calibrated to trade elasticity dynamics			
θ	Demand elasticity	3.55	ECM estimate of SR trade elasticity = 2.3
ρ_ξ	Prob. of keeping iceberg cost	0.87	ECM estimate of LR trade elasticity = 8.07
(d) Calibrated to annual NNTR gap			
$\omega_{21,t}$	Prob. NNTR to MFN	0.25	Avg. NNTR gap during 1974–1979
$\omega_{12,t}$	Prob. MFN to NNTR	Varies	NNTR gap during 1980–2008

Figure 1: China's Growing Role in the World and the U.S.

(a) China's Share of World Totals

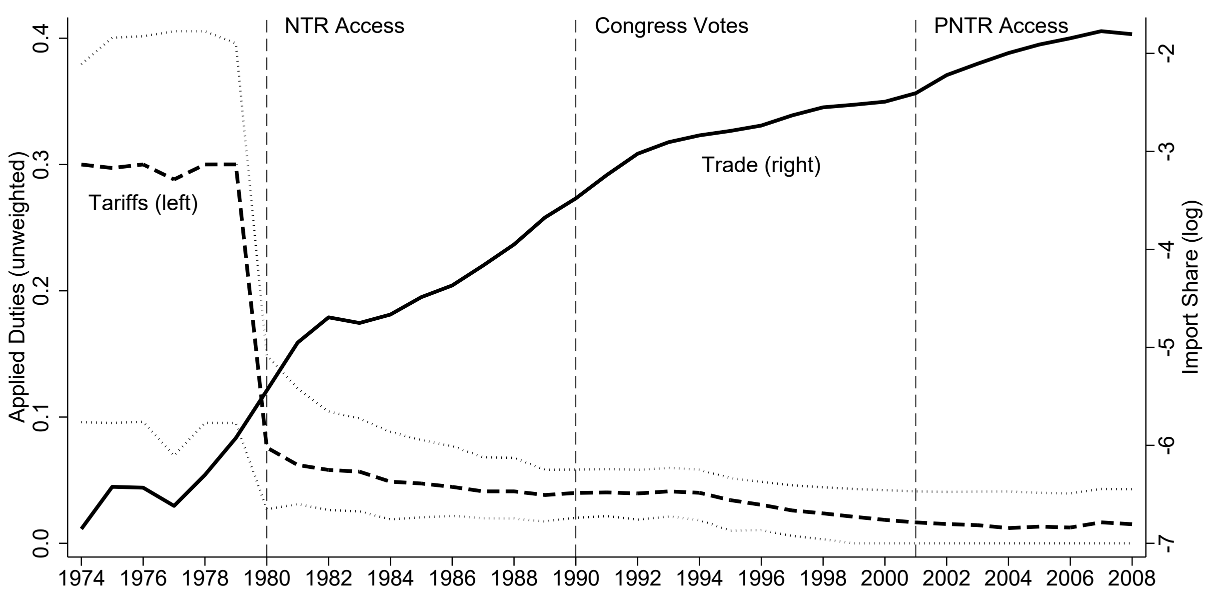


(b) China's Share of U.S. Trade



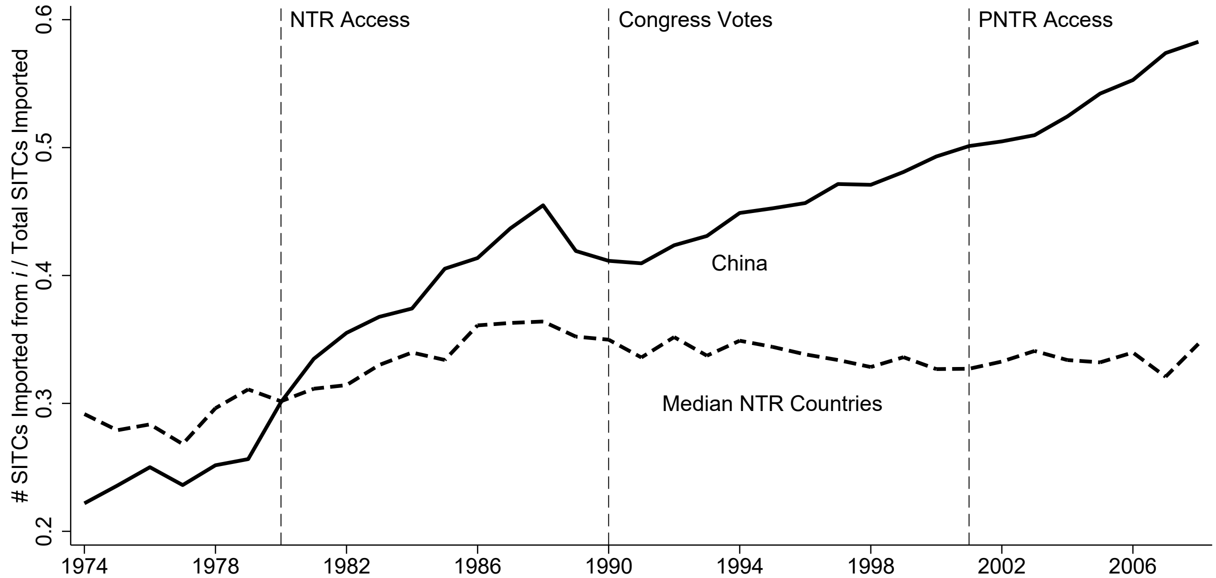
Notes: Panel (a) shows China's share of Non-U.S. World GDP (dashed) and World Exports (solid). Panel (b) shows China's share in U.S. exports (dashed) and imports (solid). Data source is UN Comtrade.

Figure 2: China's Aggregate Import Share and Duties



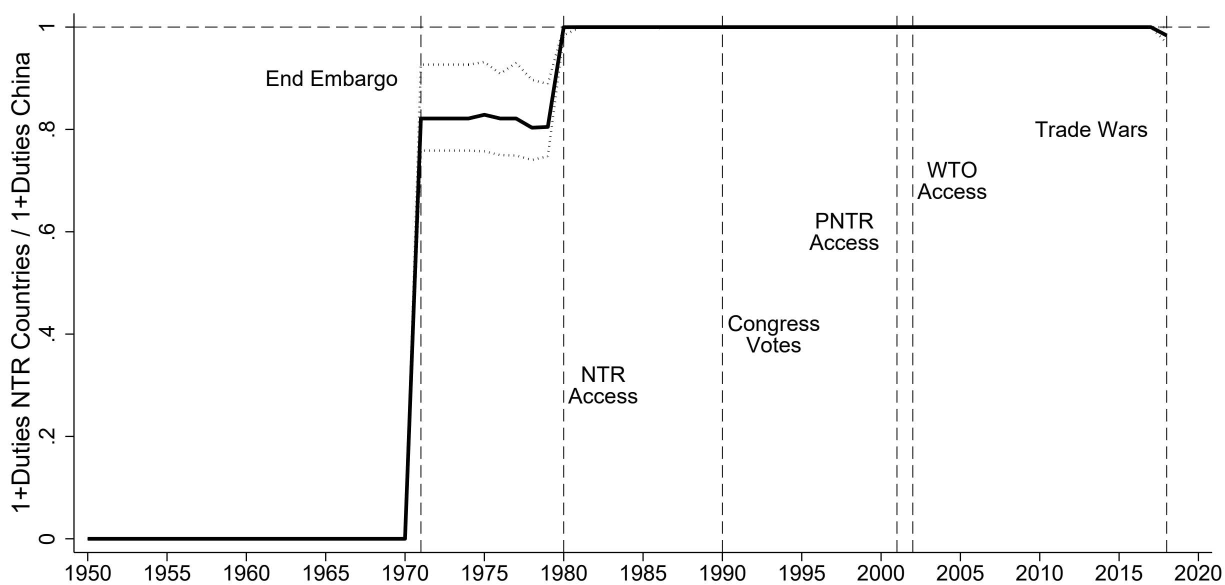
Notes: Tariffs are defined at the SITC level and calculated as the log of 1 + total duties over the total FOB value of imports. The dashed lines are the median tariff across goods and the 25th and 75th percentiles. The solid line is the log import share of China over the total of U.S. imports from all countries.

Figure 3: China's Extensive Margin of Exports to the U.S.



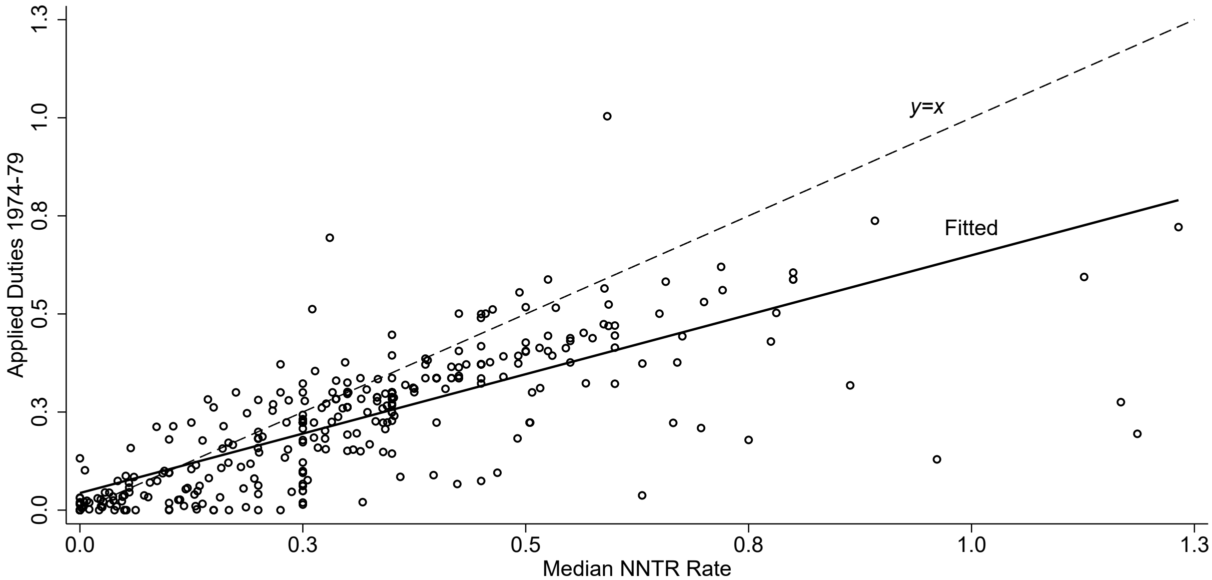
Notes: The lines are the number of imported SITC goods from China (solid) and the median of the NTR countries (dashed) over the total SITC goods imported that year. A good is considered non-traded if its share of the country's total exports to the U.S. is below 0.00001, in the spirit of Kehoe and Ruhl (2013).

Figure 4: U.S. Trade Policy toward China



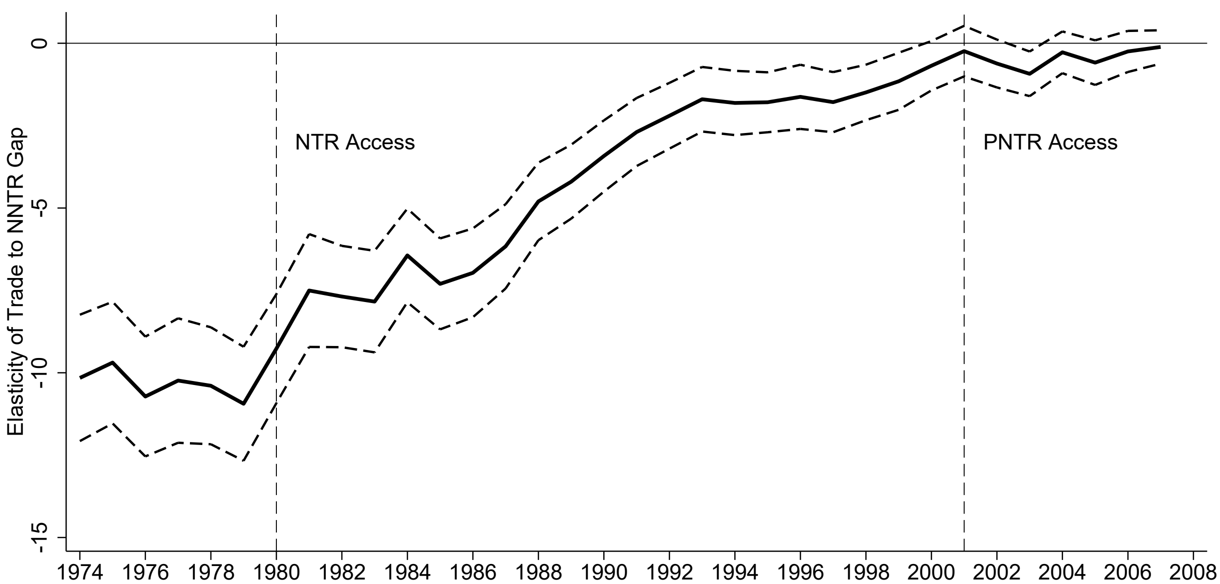
Notes: This figure plots the ratio of the 1 + tariff applied to China over the median 1 + tariff applied to NTR countries. Tariffs—total annual duties over the total annual FOB value of imports—are defined at the 8-digit TS-USA level between 1974 and 1988 and at the HS-8 level between 1989 and 2018. The solid line is the median of all products and the dashed lines are the 25$^{\text{th}}$ and 75$^{\text{th}}$ percentiles.

Figure 5: Average Applied Duties 1974–79 and the NNTR Rate



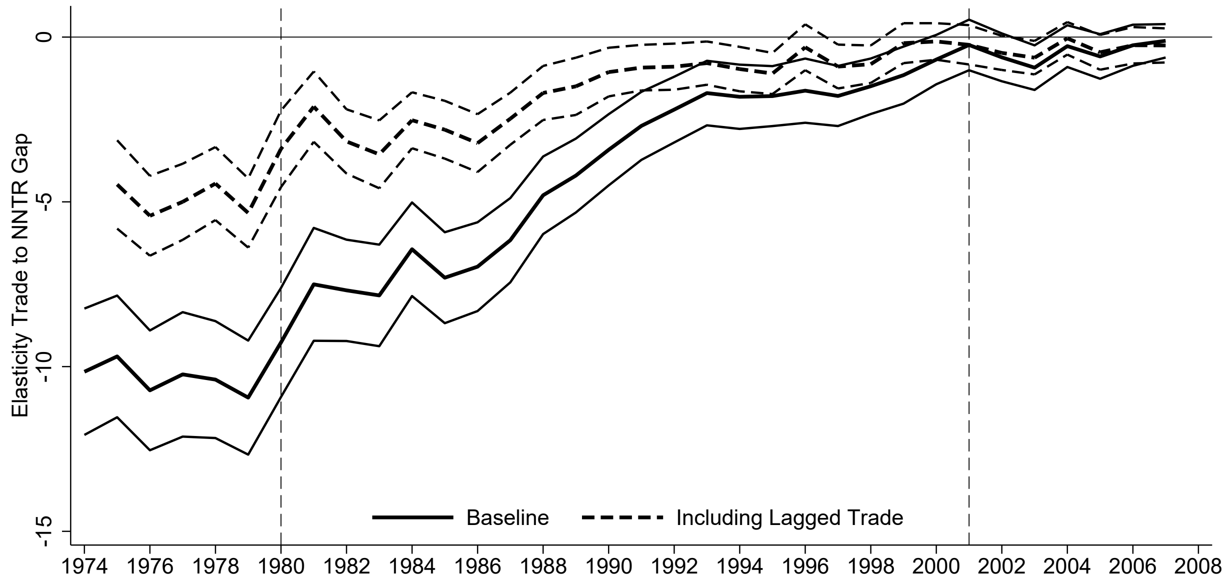
Notes: Each dot is a SITC good (excluding MFA goods). The unconditional correlation is 77 percent and the slope of the linear fit is 0.6 (solid line).

Figure 6: Elasticity of U.S. Imports from China to NNTR Gap



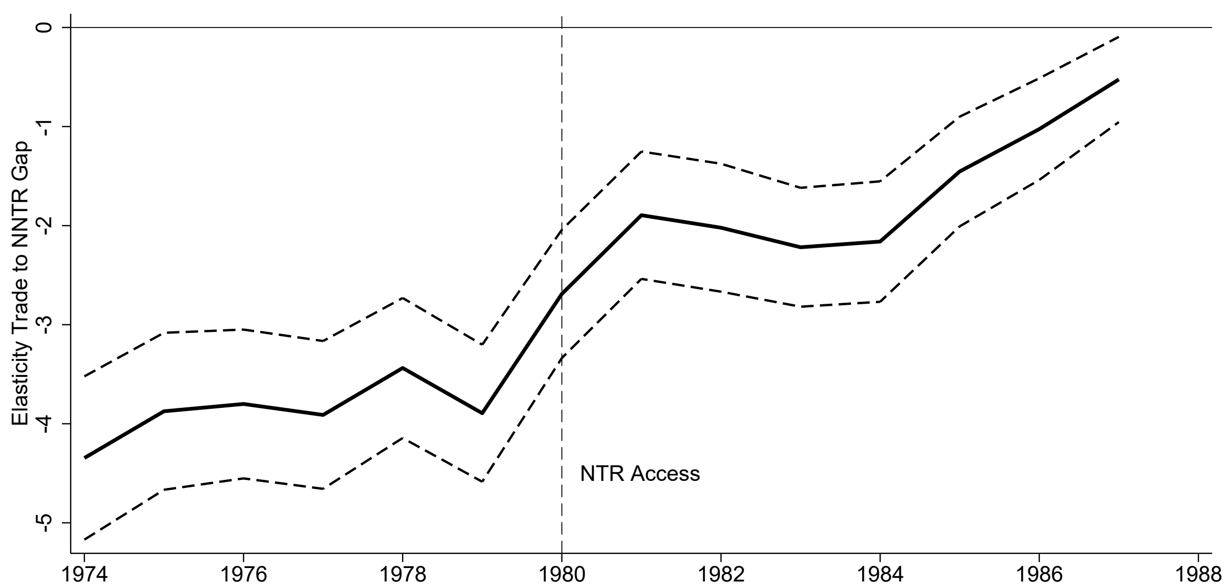
Notes: This Figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (5). The dashed lines are the 95 percent confidence interval. Standard errors are clustered at the gj level.

Figure 7: Elasticity of U.S. Imports from China to NNTR Gap, with Lagged Imports



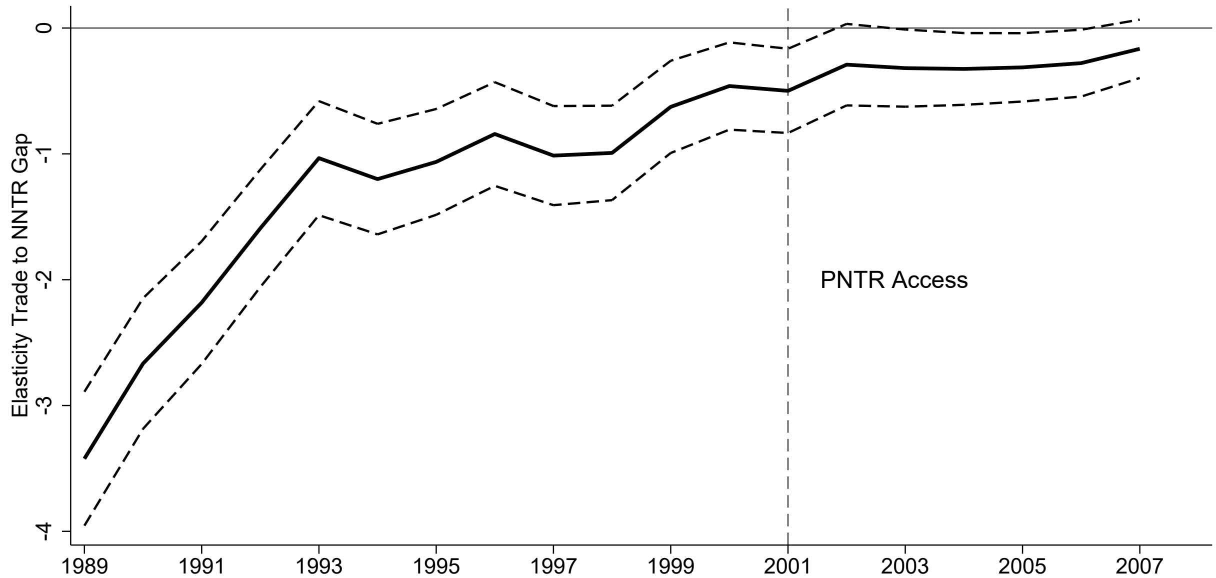
Notes: This Figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (5), including the lagged value of imports $v_{jg,t-1}$ on the right-hand side (dashed lines). The standard errors that construct the 95 percent confidence interval are clustered at the gj level.

Figure 8: Annual Elasticity to NNTR Gap: TSUSA-8 Aggregation, 1974–87



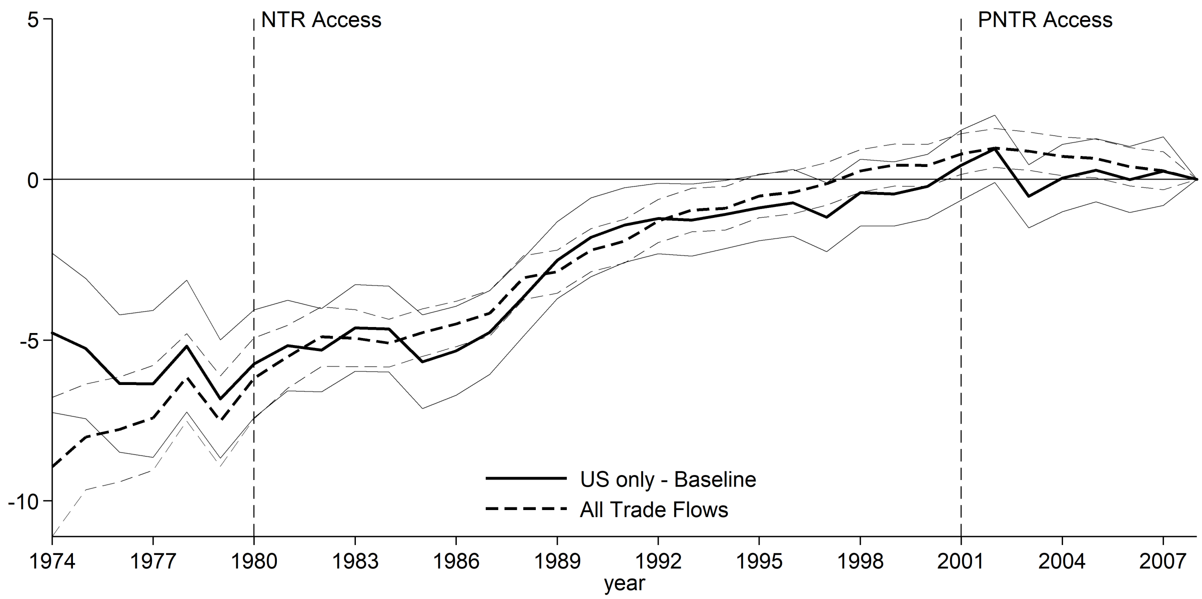
Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 1987]$ from (5) using 8-digit TSUSA product-level aggregation of all variables, instead of SITC. This level of aggregation ends in 1988. The dashed lines are the 95 percent confidence interval. Standard errors are clustered at the gj level.

Figure 9: Annual Elasticity to NNTR Gap: HS-8 Aggregation, 1989–2007



Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1989, 2007]$ from (5) using HS-8 product level aggregation of all variables, instead of SITC. This level of aggregation is only available since 1989. The dashed lines are the 95 percent confidence interval. Standard errors are clustered at the gj level.

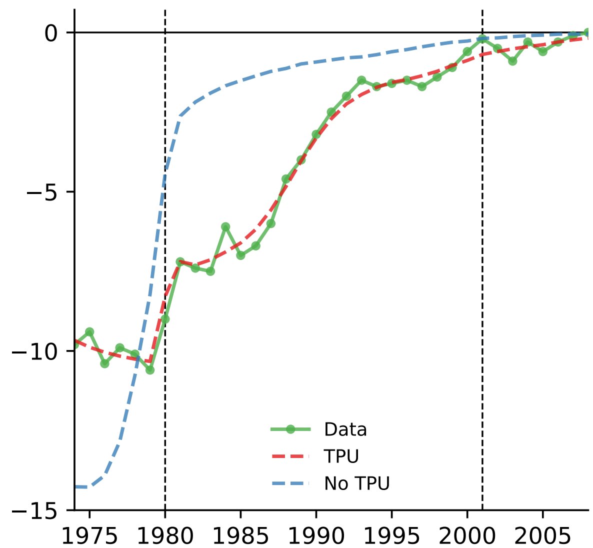
Figure 10: Annual Elasticity to NNTR Gap: China Supply Factors



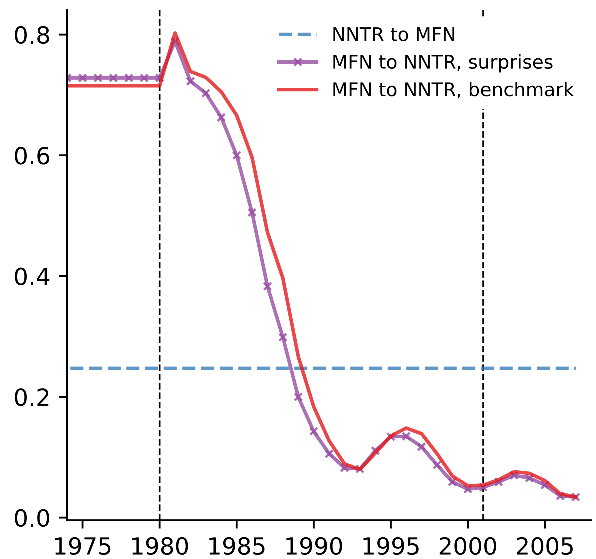
Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (6) using the merged World Trade dataset from Feenstra et al. (2005) (1974–2000) and the BACI Trade Dataset (2000–08). The dashed lines are 95 percent confidence interval. Standard errors are clustered at the gj level.

Figure 11: Trade and Policy Dynamics in the Model

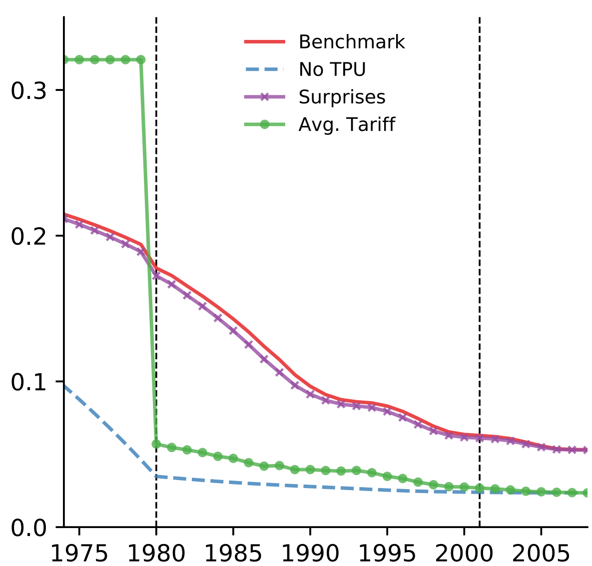
(a) Annual elasticity to NNTR gap



(b) Trade-policy probabilities



(c) Expected present value of tariffs



(d) Aggregate trade volumes (log, 1974 = 0)

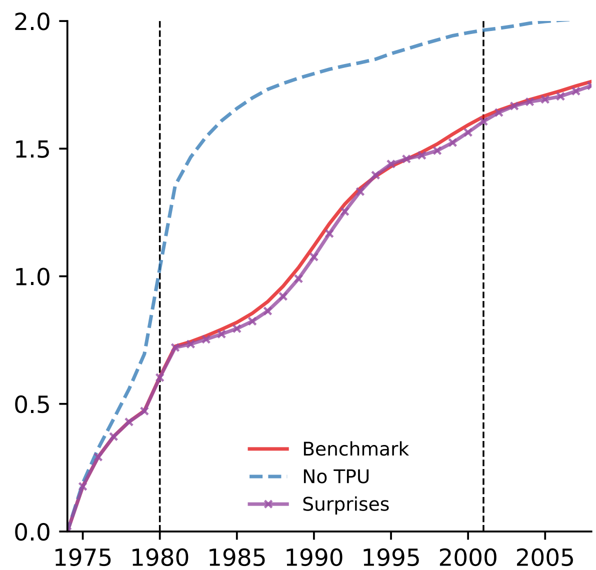


Figure 12: Model-Implied Expectations versus Bayesian Learning

(a) Prior PDF (b) Posterior mean

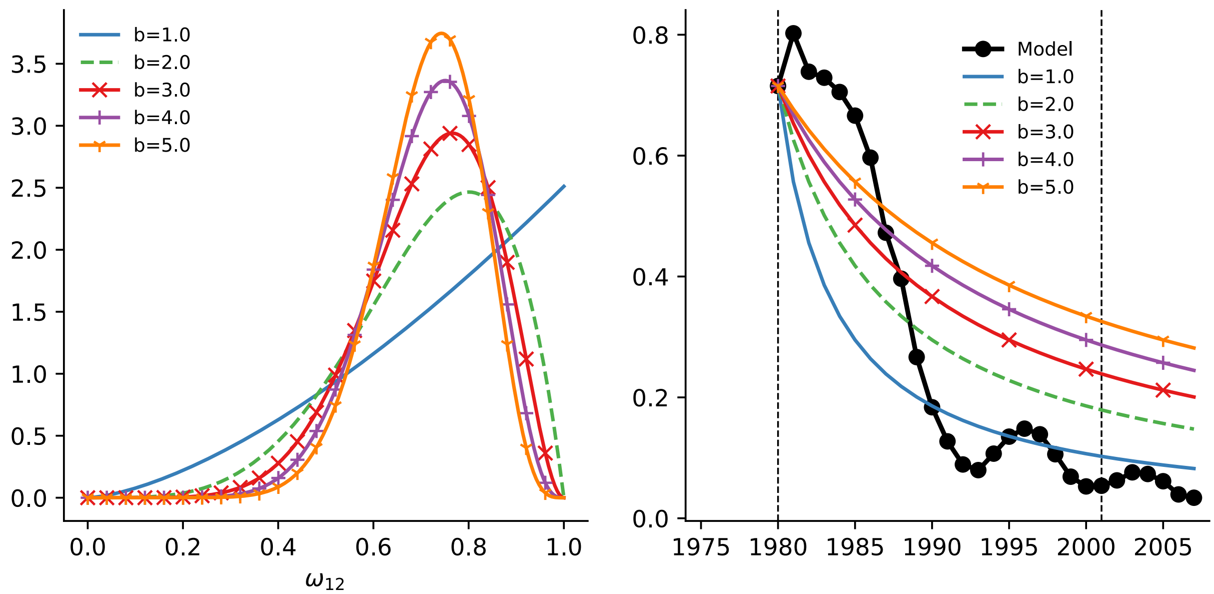
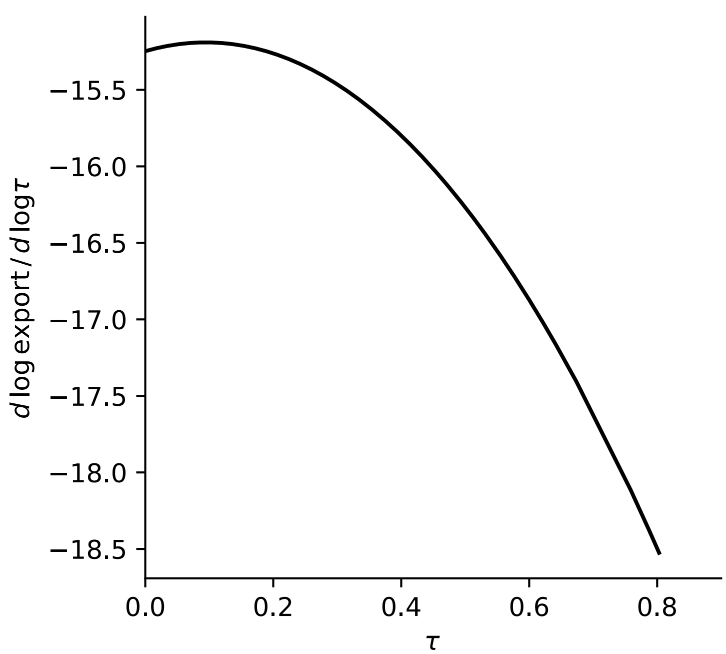


Figure 13: Long-Run Trade Elasticity in the Model



Appendix

A Sensitivity to Aggregation & Time Horizon

In section 3.3.3 we showed that the gradual effect of the NNTR gap on China's exports to the United States is also present when considering trade at the more disaggregate 8-digit TS-USA and the HS-8 level. Here we study how the level of aggregation affects the long- to short-run trade ratio. The findings indicate that (1) using a more aggregate level of trade flows (SITC) does not change the average speed of adjustment to tariffs; and (2) it is important to use the long sample period to capture the full extent of the gradualness.

Precisely, we estimate the unrestricted version of the ECM in (3) for the sample period of 1974–88 using the 8-digit TS-USA level of aggregation to compare it with the results of using the baseline SITC level. We do the same for the 1989–2008 period comparing the results of the HS-8 level of aggregation with those using SITC. Table B.3 reports the results. There are two main takeaways. First, China's long- to short-run elasticity is smaller when splitting the baseline sample period into two while continuing to use the SITC level of aggregation — compare the baseline ratio of 3.5 (column 1) to that of 1974-88 at 2 (column 3) and that of 1989-2008 at 2.7 (column 5). This is consistent with the documented slow adjustment to the 1980 NTR liberalization that extended well into the 1990s. Second, the long- to short-run elasticity ratio is not substantially affected by the level of aggregation. This is explained by the fact that, both the long- and the short-run elasticities are mostly unaffected, except for the case of NTR countries between 1989–2008 with HS-8 data (column 4), in which case both are close to two times larger in with HS-8 data. For its part, the long run elasticity is largely unaffected because the larger coefficients on the lagged tariffs ($-\alpha_2$) at the more disaggregate level are cancelled by the larger the coefficients of the lagged value of imports ($-\alpha_1$) at the more aggregate level.

B Tables

Table B.1: Annual Elasticity to NNTR Gap - TS-USA, 1974-88

Dep. Var.: v_{jgt}	Baseline Sample (1)	All Countries (2)	Applied Tariffs Communists, 1974-79 (3)	Incl. Tariffs (4)
$\mathbb{1}\{t=1974\}\mathbb{1}\{j=China\}X_g$	-4.35***	-4.23***	-4.02***	-3.01***
	(0.42)	(0.42)	(0.43)	(0.43)
$\mathbb{1}\{t=1975\}\mathbb{1}\{j=China\}X_g$	-3.87***	-3.81***	-3.55***	-2.53***
	(0.40)	(0.40)	(0.41)	(0.42)
$\mathbb{1}\{t=1976\}\mathbb{1}\{j=China\}X_g$	-3.80***	-3.69***	-3.49***	-2.43***
	(0.38)	(0.38)	(0.39)	(0.40)
$\mathbb{1}\{t=1977\}\mathbb{1}\{j=China\}X_g$	-3.91***	-3.80***	-3.59***	-2.53***
	(0.38)	(0.38)	(0.38)	(0.39)
$\mathbb{1}\{t=1978\}\mathbb{1}\{j=China\}X_g$	-3.44***	-3.27***	-3.00***	-2.09***
	(0.36)	(0.36)	(0.36)	(0.37)
$\mathbb{1}\{t=1979\}\mathbb{1}\{j=China\}X_g$	-3.89***	-3.78***	-3.54***	-2.54***
	(0.35)	(0.35)	(0.35)	(0.37)
$\mathbb{1}\{t=1980\}\mathbb{1}\{j=China\}X_g$	-2.69***	-2.61***	-2.25***	-2.48***
	(0.33)	(0.33)	(0.33)	(0.33)
$\mathbb{1}\{t=1981\}\mathbb{1}\{j=China\}X_g$	-1.89***	-1.79***	-1.78***	-1.75***
	(0.33)	(0.33)	(0.32)	(0.33)
$\mathbb{1}\{t=1982\}\mathbb{1}\{j=China\}X_g$	-2.02***	-1.93***	-1.87***	-1.97***
	(0.33)	(0.33)	(0.32)	(0.33)
$\mathbb{1}\{t=1983\}\mathbb{1}\{j=China\}X_g$	-2.22***	-2.15***	-1.93***	-2.19***
	(0.31)	(0.31)	(0.30)	(0.31)
$\mathbb{1}\{t=1984\}\mathbb{1}\{j=China\}X_g$	-2.16***	-2.05***	-1.84***	-2.14***
	(0.31)	(0.31)	(0.30)	(0.31)
$\mathbb{1}\{t=1985\}\mathbb{1}\{j=China\}X_g$	-1.45***	-1.35***	-1.09***	-1.43***
	(0.28)	(0.28)	(0.27)	(0.28)
$\mathbb{1}\{t=1986\}\mathbb{1}\{j=China\}X_g$	-1.03***	-0.93***	-0.72***	-1.00***
	(0.26)	(0.26)	(0.26)	(0.26)
$\mathbb{1}\{t=1987\}\mathbb{1}\{j=China\}X_g$	-0.52**	-0.47**	-0.32	-0.52**
	(0.22)	(0.22)	(0.22)	(0.22)
τ_{jst}				-1.97***
				(0.15)
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
Observations	455182	500641	652316	455182
Adjusted R^2	0.800	0.800	0.798	0.800

Note: All estimates are obtained by using (5), except that all variables are aggregated to 8-digit TS-USA product level, instead of the SITC level as in our baseline and that the average 1974-79 applied tariff on China is used as X_g instead of the NNTR gap (not available for TS-USA). See Figure 4 for the NTR liberalization in 1980 with TS-USA aggregation. Column 1 uses our baseline sample design that excludes NNTR and NAFTA countries as well as goods that were subject to quota removals under the MFA. Column 2 uses all countries. Column 3 uses applied tariffs to all communist countries to calculate the NNTR rate, instead of applied tariffs to China only. Column 4 includes tariffs in (5) and as expected the coefficient diminishes in the early years. Standard errors in parentheses are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table B.2: Annual Elasticity to NNTR Gap - HS-8, 1990-2007

	Baseline Sample	No Tariffs	All Countries	Full Sample	Balanced z included if $v_{CHN,g,1990} > 0$
Dep. Var.: v_{jgt}	(1)	(2)	(3)	(4)	(5)
$\mathbb{1}\{t = 1989\}\mathbb{1}\{j = China\}X_g$	-3.42***	-3.45***	-3.36***	-2.25***	-3.34***
	(0.27)	(0.27)	(0.27)	(0.23)	(0.29)
$\mathbb{1}\{t = 1990\}\mathbb{1}\{j = China\}X_g$	-2.67***	-2.70***	-2.60***	-1.82***	-2.58***
	(0.26)	(0.27)	(0.26)	(0.23)	(0.28)
$\mathbb{1}\{t = 1991\}\mathbb{1}\{j = China\}X_g$	-2.18***	-2.22***	-2.14***	-1.55***	-2.10***
	(0.25)	(0.25)	(0.25)	(0.22)	(0.27)
$\mathbb{1}\{t = 1992\}\mathbb{1}\{j = China\}X_g$	-1.64***	-1.55***	-1.09***	-1.54***	
	(0.24)	(0.24)	(0.24)	(0.20)	(0.27)
$\mathbb{1}\{t = 1993\}\mathbb{1}\{j = China\}X_g$	-1.03***	-1.08***	-0.98***	-0.83***	-1.01***
	(0.23)	(0.23)	(0.23)	(0.20)	(0.26)
$\mathbb{1}\{t = 1994\}\mathbb{1}\{j = China\}X_g$	-1.20***	-1.26***	-1.18***	-1.07***	-1.14***
	(0.22)	(0.22)	(0.22)	(0.19)	(0.26)
$\mathbb{1}\{t = 1995\}\mathbb{1}\{j = China\}X_g$	-1.06***	-1.10***	-0.99***	-1.00***	-1.21***
	(0.21)	(0.21)	(0.21)	(0.19)	(0.25)
$\mathbb{1}\{t = 1996\}\mathbb{1}\{j = China\}X_g$	-0.84***	-0.87***	-0.80***	-0.84***	-0.80***
	(0.21)	(0.21)	(0.21)	(0.18)	(0.25)
$\mathbb{1}\{t = 1997\}\mathbb{1}\{j = China\}X_g$	-1.01***	-1.03***	-0.99***	-0.98***	-1.00***
	(0.20)	(0.20)	(0.20)	(0.17)	(0.24)
$\mathbb{1}\{t = 1998\}\mathbb{1}\{j = China\}X_g$	-0.99***	-1.00***	-0.99***	-1.11***	-0.88***
	(0.19)	(0.19)	(0.19)	(0.17)	(0.22)
$\mathbb{1}\{t = 1999\}\mathbb{1}\{j = China\}X_g$	-0.63***	-0.63***	-0.62***	-0.89***	-0.57**
	(0.19)	(0.19)	(0.19)	(0.16)	(0.23)
$\mathbb{1}\{t = 2000\}\mathbb{1}\{j = China\}X_g$	-0.46***	-0.47***	-0.48***	-0.88***	-0.31
	(0.18)	(0.18)	(0.18)	(0.16)	(0.21)
$\mathbb{1}\{t = 2001\}\mathbb{1}\{j = China\}X_g$	-0.50***	-0.52***	-0.51***	-0.99***	-0.29
	(0.17)	(0.17)	(0.17)	(0.15)	(0.20)
$\mathbb{1}\{t = 2002\}\mathbb{1}\{j = China\}X_g$	-0.29*	-0.30*	-0.30*	-0.60***	-0.15
	(0.17)	(0.16)	(0.16)	(0.14)	(0.20)
$\mathbb{1}\{t = 2003\}\mathbb{1}\{j = China\}X_g$	-0.32**	-0.32**	-0.35**	-0.56***	-0.33*
	(0.16)	(0.16)	(0.16)	(0.14)	(0.19)
$\mathbb{1}\{t = 2004\}\mathbb{1}\{j = China\}X_g$	-0.33**	-0.32**	-0.32**	-0.54***	-0.18
	(0.15)	(0.15)	(0.14)	(0.13)	(0.17)
$\mathbb{1}\{t = 2005\}\mathbb{1}\{j = China\}X_g$	-0.31**	-0.31**	-0.31**	-0.069	-0.086
	(0.14)	(0.14)	(0.14)	(0.12)	(0.16)
$\mathbb{1}\{t = 2006\}\mathbb{1}\{j = China\}X_g$	-0.28**	-0.28**	-0.31**	-0.14	-0.18
	(0.14)	(0.14)	(0.14)	(0.11)	(0.15)
$\mathbb{1}\{t = 2007\}\mathbb{1}\{j = China\}X_g$	-0.17	-0.16	-0.19	-0.023	0.050
	(0.12)	(0.12)	(0.12)	(0.10)	(0.13)
τ_{jst}	-3.87***		-3.37***	-4.07***	-4.62***
	(0.20)		(0.26)	(0.22)	(0.19)
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
Observations	2079150	2079150	2314918	2930184	1246200
Adjusted R^2	0.766	0.766	0.774	0.774	0.778

Note: All estimates are obtained by using (5), except and that all variables are aggregated to HS-8 product level, instead of the SITC level as in our baseline, and that τ_{jgt} is added to the right hand side of (5) (as in Pierce and Schott (2016)). Column 1 uses our baseline sample design that excludes NNTR and NAFTA countries as well as goods that were subject to quota removals under the MFA. Column 2 uses the same sample but excludes tariffs from the regression. Column 3 uses all countries and column 4 further includes all goods. Column 5 uses only products from which U.S. imports from China was non-zero in 1990. Standard errors in parentheses are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table B.3: Sensitivity of Long- & Short-Run Elasticities to Aggregation & Time Horizon

ECM Unrestricted	1974-2008	1974-88		1989-2008	
	SITC	TS-USA	SITC	HS-8	SITC
Dep. Var.: Δv_{jzt}	(1)	(2)	(3)	(4)	(5)
$\mathbb{1}\{j \neq China\}\Delta\tau_{jst}$	-2.02***	-2.24***	-2.23***	-3.79***	-1.76***
	(0.24)	(0.13)	(0.39)	(0.13)	(0.29)
$\mathbb{1}\{j = China\}\Delta\tau_{jst}$	-2.30***	-1.89***	-2.33***	-2.26***	-2.21**
	(0.41)	(0.28)	(0.38)	(0.57)	(0.97)
$\mathbb{1}\{j \neq China\}v_{js,t-1}$	-0.47***	-0.78***	-0.67***	-0.60***	-0.54***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$\mathbb{1}\{j = China\}v_{js,t-1}$	-0.37***	-0.75***	-0.61***	-0.51***	-0.45***
	(0.01)	(0.01)	(0.02)	(0.00)	(0.01)
$\mathbb{1}\{j \neq China\}\tau_{js,t-1}$	-1.61***	-2.76***	-1.89***	-3.27***	-1.74***
	(0.16)	(0.16)	(0.29)	(0.12)	(0.22)
$\mathbb{1}\{j = China\}\tau_{js,t-1}$	-2.96***	-3.11***	-2.86***	-3.88***	-2.72***
	(0.22)	(0.27)	(0.26)	(0.40)	(0.68)
Long-Run NTR Countries	-3.45***	-3.51***	-2.83***	-5.44***	-3.21***
	(0.35)	(0.20)	(0.43)	(0.20)	(0.41)
Long-Run China	-8.07***	-4.15***	-4.66***	-7.55***	-6.06***
	(0.57)	(0.36)	(0.41)	(0.78)	(1.51)
Long-/Short-Run NTR Countries	1.71	1.57	1.27	1.43	1.82
Long-/Short-Run China	3.51	2.20	2.00	3.34	2.74
FE	st, jt, sj	st, jt, sj	st, jt, sj	st, jt, sj	st, jt, sj
Observations	934249	882590	297073	2122714	631559
Adjusted R^2	0.269	0.383	0.361	0.291	0.292

Note: All estimates are obtained by using (3). Column 1 uses the baseline sample period (1974–2008) and aggregation level of variables (5-digit SITC). Columns 2 and 3 use the sample period 1974–89, and columns 4 and 5 1989–2008. Column 2 uses 8-digit TS-USA aggregation of all variables. Column 3 and 5 SITC and column 4 8-digit HS level. Fixed effects of s are at the aggregation level of the sample used. Standard errors in parentheses are clustered at the zj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.