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

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Abstract

We study the growth of Chinese exports to the United States, from autarky during 1950–1970 to 15 percent of overall U.S. imports in 2008, taking advantage of the rich heterogeneity in trade policy and trade growth across products during this period. Central to our analysis is an accounting for the dynamics of trade flows, observed trade policy, and expectations about future policy. In our empirical analysis, we estimate the dynamics of the elasticity of Chinese exports to (i) past tariff changes and (ii) the risk of future tariff hikes. We find that Chinese exports responded slowly to the tariff changes that occurred when China was granted Most Favored Nation status in 1980, and that policy uncertainty was more important in the immediate aftermath of this liberalization than in the lead-up to China's 2001 accession to the World Trade Organization. It is difficult, however, to separately identify these two effects using data alone. In our quantitative analysis, we disentangle these effects by using a structural model to estimate a path of trade-policy expectations. We find that the 1980 reform was largely a surprise and initially had a high probability of being reversed. The likelihood of reversal dropped considerably during the mid 1980s but changed little throughout the late 1990s and early 2000s despite China's accession to the World Trade Organization in 2001.

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1 Introduction

International trade depends on past, present, and future trade policy, but rarely are the three studied together. 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 changes in trade policy that occurred in the past or may occur in the future.[1] One recent literature shows that trade responds gradually to past policy changes, while another shows that uncertainty about future policy can affect trade in the present. We argue that the effects of past and future policy are tightly intertwined, develop a method to disentangle these effects, and apply it to Chinese exports to the United States from 1971 to 2008.

We find that accounting for innovations to both past and future trade policies substantially changes our understanding of the timing of policy changes and their influence on the dynamics of Chinese exports. We argue for a smaller role for policy uncertainty during the 1990s in explaining these dynamics than other studies have estimated, and a larger role in the early 1980s. We find that Chinese exports were depressed in the early 1980s after China was granted Most-Favored-Nation (MFN) status because this status was expected to be revoked. This reform became more credible in the late 1980s and early 1990s, and this change in expectations was the main driver of the sustained growth of Chinese exports during the lead-up to, and wake of, China's 2001 accession to the World Trade Organization (WTO). Contrary to other studies, we find that WTO accession itself had only a small impact on expectations about U.S. trade policy and Chinese export growth. Thus, we offer a substantially different narrative about the role of changes in trade policy, past and future, on the growth of U.S. trade with China.

China's integration with the U.S. economy is ideally suited to the study of trade-policy dynamics. China started from autarky, as the United States maintained a complete embargo on Chinese goods from 1950 until 1971. From autarky, two major trade liberal-

[1]The trade elasticity is generally measured in the cross-section and identified by the effect of cross-country differences in current tariffs on cross-country differences in current 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.

izations occurred in 1971 and 1980. The first granted Non-Normalized Trade Relations (NNTR) status to China, allowing Chinese goods to enter the U.S. market at relatively high tariff rates. The second granted Normal Trade Relations (NTR) status to China, lowering tariffs on U.S. Chinese goods to MFN levels. These two liberalizations were large, immediate, and heterogeneous across goods, providing the cross-industry variation that is key to our identification of the trade response. The duration of these liberalizations, however, was uncertain. The program to transition communist countries from NNTR to MFN tariff rates, which began in the mid 1970s, required the United States to renew China's NTR status annually. The renewal process applied to all products, but the potential effect of non-renewal was heterogeneous across industries, owing to differences in the gap between NNTR and MFN tariffs. When China joined the WTO and gained permanent NTR status in 2001, the United States eliminated this renewal process, potentially reducing uncertainty about U.S. tariffs on Chinese goods.[2] In total, these reforms propelled China from the United States' smallest trading partner to its largest.

Our methodology requires two empirical measurements that are inputs into a structural model. The first is a measure of the gradual adjustment of trade to a change in tariffs. Our dataset covers the dynamics of Chinese exports to the United States following the 1980 tariff reform and the ensuing two decades. Using an error-correction model (Johnson and Oksanen, 1977; Johnson et al., 1992; Gallaway et al., 2003) and a local-projections specification (Jordà, 2005; Boehm et al., 2020), we find that the long-run response of trade to tariffs is almost four times as large as the short-run response and that it takes as long as 20 years to complete 90 percent of the adjustment.

The second empirical measurement estimates the effect of tariff risk on Chinese exports. Building on Pierce and Schott (2016), we measure the response of trade in goods with large differences between NNTR and MFN tariffs relative to the response of goods with small differences. These tariff "gaps" measure the increase in tariffs goods would face if China lost NTR status. If policy uncertainty is important, trade in high-gap goods will be lower than low-gap goods. We measure the elasticity of Chinese exports to the

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

tariff gap for each year from 1974–2007. We find the effects of policy uncertainty to be the largest in the 1970s and early 1980s, and that these effects had largely dissipated before China joined the WTO in 2001.[3]

By themselves, neither of these measurements identify the effects of past policy or future policy risk because the NNTR gap is almost perfectly correlated with the change in tariffs in 1980. The long- and short-run elasticities we measure are biased by trade policy uncertainty (TPU). If the risk of losing NTR status depressed trade prior to China's WTO accession, then the adjustment process we measure is slower than it would have been in the absence of this uncertainty. The tariff-gap elasticities are biased by the slow adjustment from the original liberalization. Goods with the largest gaps between NNTR and MFN tariffs are most sensitive to policy uncertainty, but these goods also experienced the largest tariff reductions in 1980 and will take the longest to converge to their new level of trade. Our methodology allows us to overcome these problems by estimating the time-varying transition probabilities between these two policy regimes using a dynamic exporting model.

Our model is a multi-industry version of the heterogeneous-firm model with sunk export costs and new exporter dynamics developed by Alessandria et al. (2021a). This is a generalization of the sunk-cost exporting model of Baldwin (1988), Dixit (1989), and Das et al. (2007) that captures the key features of marginal exporter dynamics as emphasized by Ruhl and Willis (2017). Industries in the model correspond directly to the goods used in our empirical analysis. Firms in each industry differ in terms of productivity, as well as variable export costs which they reduce gradually through a risky investment.[4] The export entry decision and gradual reduction in export costs cause trade volumes to adjust slowly to changes in tariffs. Thus, past policy can affect trade long after its implementation. The model features two trade-policy regimes, NNTR and MFN, and the probability of switching between regimes varies over time. As in Handley and Limão (2017), this uncertainty depresses export participation and reduces trade volumes when the economy is

[3]The fact that the effect of the NNTR gap on trade was largest during the 1970s, before China gained NTR status in the first place, suggests that the NNTR gap may capture something else in addition to exposure to tariff risk. This is precisely one of the points of our paper.

[4]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.

in the MFN regime.

The model is calibrated using indirect inference so that it matches our estimates of the adjustment process and the elasticities of trade to the NNTR gap. Firms in the model understand that the tariff regime can change in any period, but the realized path of trade policy is identical to the historical experience: the model begins in 1971 in the NNTR regime and switches to the MFN regime in 1980. The probabilities of switching between trade policy regimes are chosen so that the transition to the steady state replicates our estimated time path of the elasticities of trade to the NNTR gap. Thus, we 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 exports of the former relative to the latter.

We estimate a time-varying path of transition probabilities between MFN and NNTR status over a longer time frame than previous work using a richer model of trade adjustment dynamics. We find the annual probability of China gaining NTR status during the 1970s was about 25 percent. Once China gained NTR status in 1980, the probability of losing this status was initially high, peaking at 81 percent in 1981. This 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, the probability of losing NTR status began to fall rapidly. It temporarily rose again in the early-to-mid 1990s in the wake of the Tiannanmen Square incident, but by the late 1990s it had fallen to 5–10 percent. Joining the WTO had a minor effect on the probability of losing access to MFN tariff rates; this probability fell by less than 2 percentage points between 1999 and 2008.

Our model allows us to measure the effect of alternative expectations of future tariff policy on trade, while controlling for the slow adjustment that biases the empirical elasticities. When there is perfect foresight over trade policy, so that the changes are viewed as permanent, 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 actual estimated value a few years after the 1980 liberalization.

Our model also highlights the importance of a subtle yet important aspect of trade adjustment dynamics: trade adjusts slowly to changes in expectations about future policy as well as policy changes that have occurred in the past. To illustrate this point, we ask our model: What would have happened if the credibility of China's NTR status had not grown steadily throughout the 1980s and 1990s, and had instead remained constant until China joined the WTO? In this experiment, we calibrate our model so that it matches Pierce and Schott (2016)'s estimate of how the elasticity of Chinese exports to the NNTR gap changed after 2001, rather than matching our estimated time series for the annual NNTR-gap elasticity. This increases the probability of losing NTR status in the late 1990s by more than 20 percentage points, significantly overstating the importance of WTO accession on tariff risk. This demonstrates that time-varying policy uncertainty is a key factor in explaining the path of China's integration into the U.S. market. Specifically, much of the growth in exports of goods with high NNTR gaps after the turn of the century was a delayed adjustment to the increase in the credibility of U.S. policy towards China during the 1980s and 1990s, rather than a reduction in tariff risk associated with WTO accession and PNTR status.

This paper is related to two main strands of literature. The first strand studies the dynamics of trade flows after changes in trade policy. Baier and Bergstrand (2007) and Baier et al. (2014) show that trade grows slowly following the creation of a free trade area, with only one-third of the long-run response occurring in the first few years. Similarly, Khan and Khederlarian (2021), Anderson and Yotov (2020), Boehm et al. (2020) estimate long-run tariff elasticities of trade that are three to four times short-run elasticities. Many of these studies grapple with issues related to endogeneity and anticipation of phased-in and/or temporary tariff changes. Our contribution is to study the dynamic response to a single exogenous, immediate tariff reduction over a multi-decade time span using disaggregated good-level data. We document similar results using two entirely different empirical specifications: an error-correction model that recovers short- and long-run trade elasticities while imposing a parametric adjustment path (Johnson and Oksanen, 1977; Johnson et al., 1992; Gallaway et al., 2003); and a local-projections model that recovers a non-parametric impulse response (Jordà, 2005; Boehm et al., 2020).

The second strand studies how trade reforms that are expected to occur in the future affect trade in the present. 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.[5] Pierce and Schott (2016), Feng et al. (2017), and Handley and Limão (2017) leverage the well-defined nature of this uncertainty—the annual vote to retain China's NTR status and the gap between the MFN tariffs and the fallback NNTR tariffs—to measure the growth in trade that resulted from the elimination of uncertainty that occurred when China was granted Permanent Normal Trade Relations (PNTR) status in 2000. Similarly, Bianconi et al. (2021) study the effect of the NNTR gap on U.S. firms' stock returns. Our contribution is to study how these tariff gaps, which also capture the size of the initial 1980 liberalization, influenced trade dynamics from the beginning of the U.S.-China trade relationship. We find that studying the reforms over a longer window helps to identifying trade policy uncertainty.

Our quantitative analysis, which disentangles the effects of slow adjustment and tariff risk on trade, contributes to both of these strands of literature. Several other papers use models to estimate trade policy expectations. Early work by Ruhl (2011) estimates the probability of ending the ban on Canadian beef after an outbreak of "mad cow disease." Like us, Handley and Limão (2017) use a dynamic exporting model and the evidence on differential trade growth across products to estimate the probability of NNTR reversal. We estimate a time-varying probability over a longer interval using a richer model of ex-

[5]There are also several papers that study the impact of uncertainty about Brexit, such as Graziano et al. (2018), Steinberg (2019), and Crowley et al. (2019).

porter dynamics in which changes in tariffs have highly persistent effects. Our analysis highlights the importance of earlier changes in trade-policy expectations during the late 1980s and early 1990s to explaining the growth of Chinese exports to the United States around the turn of the century; we find that ignoring these changes leads one to overstate the effect of PNTR access on tariff risk. Complementary to our approach, Alessandria et al. (2019) estimate a time-varying probability of NNTR reversal from 1990 to 2005 using within-year variation in trade flows and trade-policy risk in an sS inventory model, obtaining similar results for this later period. Our work is also related to Alessandria et al. (2017), who estimate the expected path of inward and outward trade policy for China from macroeconomic time series.

Section 2 describes our dataset, presents the results of our empirical analyses, and discusses why these analyses cannot separate the effect of slow adjustment from the effect of tariff risk. Section 3 lays out our model and discusses our calibration strategy. Section 4 presents the results of our quantitative analysis. Section 5 concludes.

2 An Empirical History of U.S.-China Trade

We take a two-pronged approach to empirically analyzing the joint dynamics of U.S. trade policy towards China and imports of Chinese goods. First, building on the trade adjustment dynamics literature, we estimate short- and long-run elasticities of trade to the 1980 tariff reduction and the speed of adjustment between these two horizons. Second, building on the trade-policy uncertainty literature, we study the elasticity of trade to the risk of reversing this tariff reduction, and how this elasticity changed over time. Although neither of these approaches can disentangle the effects of slow adjustment to past policy changes from uncertainty about future policy or identify structural parameters, they both produce crucial inputs to our quantitative analysis.

2.1 Data

We use annual data on U.S. imports from 1974 to 2008, aggregated at the 5-digit level of the Standard International Trade Classification (SITC), revision 2. This level of aggregation provides continuous coverage of almost the entire history of U.S.-China trade.[6] We refer to this level of aggregation as a good and denote it by g. There are 1,742 goods in our sample. Our import data include applied duties, cost-insurance-and-freight (CIF) charges, and the free-on-board (FOB) import value. The log FOB import value is denoted by v_{gjt}, where j indexes the exporting country and t indexes time.

We use two measures of trade policy: applied and statutory tariff rates. We calculate applied tariffs, denoted τ_{jgt}, by dividing applied duties by FOB import values. We obtain ad-valorem-equivalent NNTR and MFN statutory tariffs from Feenstra et al. (2002) at the 8-digit level of the Harmonized Tariff Schedule (HS), and then map them to our 5-digit SITC classification using the concordance from United Nations Trade Statistics (2017). The SITC-level NNTR and MFN tariffs (τ_g^{NNTR} and τ_{gt}^{MFN}, respectively) are the median 8-digit product-level tariffs within each SITC good. Both statutory tariff schedules are exogenous to China's growth and trade integration. NNTR tariffs were established by the Smoot-Hawley Act of 1930, long before the United States began trading with China, and MFN tariffs apply to all WTO members (and non-members that have been granted conditional NTR status).

Our baseline sample includes imports from China and all other countries that had NTR status throughout the entire period and were not part of preferential trade agreements with the United States. This excludes Canada, Mexico, and other communist countries.[7] A key feature of our baseline sample is that all countries, including China, faced approximately the same tariffs after 1980.[8] Our strategy is to treat non-China NTR countries as the control group; we analyze how tariffs on Chinese goods changed relative to tariffs on other NTR countries' goods, and how imports of Chinese goods grew relative to imports from other NTR countries. We also exclude goods that include products covered by the Multi Fiber Arrangement (MFA). As documented by Bambrilla et al. (2010),

[6] More disaggregated product-level data, such as 8-digit TS-USA or HS classifications, cover only portions of this period due to a change in classification schemes in 1989. As we show in the online appendix, our main results hold when using TS-USA and HS data during the periods in which these data are available.

[7] 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.

[8] Variation in bilateral applied tariffs among NTR countries is due to aggregation, specific tariffs, temporary commercial policy, or measurement error. As Figure 2b shows, however, this variation is minor.

China's accession to the WTO triggered the removal of import quotas on these goods, and this material change in applied trade policy is not captured by our tariff measures. In the appendix we show that our results are robust to alternative sample designs.

2.2 Policy dynamics

There are three key reforms in the history of U.S.-China trade: when the United States lifted its longstanding embargo on Chinese goods in 1971; when the United States granted NTR status to China in 1980; and when China joined the WTO and its NTR status was made permanent in 2001.

The 1971 and 1980 reforms changed trade policy dramatically. Before 1971, imports from China faced effectively infinite tariffs. Between 1971 and 1979, Chinese imports were taxed at the relatively high NNTR tariff rates set by the Smoot-Hawley Act of 1930. From 1980 until the 2018 trade war, Chinese goods were taxed at the much lower MFN tariffs that apply to imports from WTO members (and other non-members that, like China, have been unilaterally granted NTR status). Table 1 reports summary statistics about the NNTR and MFN tariff schedules at the 1-digit SITC level. The mean NNTR rate is 28 percent with a standard deviation of 18 percent. The average applied tariff for non-China NTR countries was five in 1979 and two in 2001. For China, the average applied rate was considerably higher in 1979 (20 versus five), but quite similar by 2001 (three versus two). Figure 1 shows how the distribution of applied tariffs on Chinese goods, summarized by the median tariff at the 5-digit SITC level and the interquartile range, changed over our sample period. The median tariff fell from 30 percent to about 8 percent. The vast majority of this decline occurred in 1980. Subsequent tariff reductions in MFN tariffs were related to gradual 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).

Figure 2 shows the evolution of U.S. trade policy towards China relative to countries that had NTR status throughout our sample period. Panel (a) plots the inverse tariff on Chinese goods normalized by the tariff rate applied to other NTR countries:

$$\tau^{\mathrm{INV}}_{China,gt} = \frac{1+\tau^{MFN}_{gt}}{1+\tau_{China,gt}}. \tag{1}$$

In 1971, when the embargo ended, the inverse relative tariff jumped from zero to 80 percent, and then in 1980, when China gained NTR status, it jumped to one. Unlike the GATT rounds, which featured gradual tariff phaseouts that were agreed upon in advance, these two relative tariff cuts were both immediate. Panel (b) plots the distribution of the residuals obtained by regressing annual tariff changes on country-year, good-year, and country-good fixed effects. Virtually all the variation in tariff changes over and above multilateral changes in MFN tariffs, which are absorbed by the fixed effects, was due to the 1980 NTR status grant. There is some small variation from 1981 onward due to specific tariffs and temporary commercial policy (as well as aggregation issues and measurement error), but this variation is minor compared to the effects of NTR access.

In addition to these changes in applied tariffs, there is another important aspect to the dynamics of U.S. trade policy towards China: uncertainty. China was granted NTR status in 1980 as part of a new program to allow non-market economies to access U.S. markets, and this grant was contingent on satisfying conditions of the Jackson-Vanik Amendment to the Trade Act of 1974.[9] The U.S. President had to renew China's NTR status each July, and each year from 1990 onward the U.S. Congress convened to vote on whether to disapprove this renewal.[10] The U.S. House of Representatives voted to revoke China's NTR status in 1990, 1991, and 1992, although the Senate did not. 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 status.

2.3 Slow adjustment to the 1980 NTR status grant

We begin our empirical analysis by studying how U.S. imports from China adjusted after the 1980 grant of NTR status. It is well known that trade adjusts gradually to trade liberalizations. For example, Baier and Bergstrand (2007) estimate that trade doubles in

[9] 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 NTR status from 1988 to 1991.

[10] The change in administration in 1981 likely also increased uncertainty about whether China's NTR status would be revoked, as President Ronald Reagan adopted more protectionist policies and was more openly anti-communist than President Carter.

the long run after the creation of a free trade area, but only around one third of this response occurs on impact. Other studies, such as Anderson and Yotov (2020), Khan and Khederlarian (2021), and Boehm et al. (2020), find similar differences between short- and long-run trade responses. We use two approaches to study the dynamic response of U.S. imports from China to the 1980 reform: an error correction model (ECM), which recovers short- and long-run trade elasticities while imposing a parametric path of adjustment (Johnson and Oksanen, 1977; Johnson et al., 1992; Gallaway et al., 2003) and local projections, which recover non-parametric impulse responses (Jordà, 2005). Both approaches show that the growth in trade that followed the 1980 reform was gradual and took many years to complete.

Our first approach is an unrestricted ECM specification:

$$\begin{aligned}\Delta v_{jgt} = &\left[\sigma^{SR}_{China}\Delta\tau_{jgt} + \gamma_{China}\left(v_{jg,t-1} - \sigma^{LR}_{China}\tau_{jg,t-1}\right)\right]\mathbb{1}_{\{j=China\}}\\ &+\left[\sigma^{SR}_{Others}\Delta\tau_{jgt} + \gamma_{Others}\left(v_{jg,t-1} - \sigma^{LR}_{Others}\tau_{jg,t-1}\right)\right]\mathbb{1}_{\{j=Others\}}\\ &+\delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}. \end{aligned} \tag{2}$$

The dependent variable is the one-year log difference in import value. The right-hand side includes the one-year change in applied tariffs, lagged tariffs, lagged log imports, and a set of fixed effects. We interact the first three variables with an indicator variable for China to estimate China-specific elasticities. The short-run trade elasticity, σ^{SR}, is the coefficient on the one-year change in tariffs. The long-run elasticity, σ^{LR}, is pinned down by the response to lagged tariffs and the autocorrelation of imports. Country-year (δ_{jt}) fixed effects capture aggregate shocks to exporting countries; country-good fixed effects (δ_{jg}) reflect the average level of exports as well as time-invariant bilateral trade barriers; and good-year fixed effects (δ_{gt}) capture good-level U.S. demand shocks as well as good-specific trade barriers common to all exporters. This set of fixed effects is standard in the empirical literature on trade dynamics with panel data.[11] Because the baseline sample excludes countries other than China that did not receive NTR status at some point in the sample, the good-year fixed effects absorb the effects of multilateral changes in MFN tariff

[11]See, for example, Head and Ries (2001) and Romalis (2007)

rates on trade; σ^{SR} and σ^{LR} measure how U.S. imports responded to changes in bilateral tariffs above and beyond these multilateral changes.

The solid line in panel (a) of Figure 3 shows the path of adjustment to a one-time tariff change implied by our ECM estimates; the estimated parameter values are reported in Table E.1 of the appendix. The short-run trade elasticity, σ^{SR}, is -2.29 and the long-run elasticity, σ^{LR}, is -7.96. The former is consistent with many other estimates in the literature, and the latter, while large, is similar to the effects of other major liberalizations.[12] The large gap between the short- and long-run responses indicates that the adjustment of U.S. imports of Chinese goods to tariff changes has been very gradual. As can be seen in panel (b), our estimates imply that it takes seven years for U.S. imports from China to complete 90 percent of the total long-run adjustment to a tariff change. We show in appendix A that these results are robust to a range of alternative specifications with additional controls and different samples of countries and goods.

Our second approach is a local projection specification:

$$\begin{aligned}\Delta_h v_{jg,1979} = {} & \sigma^h_{China} \mathbb{1}_{\{j=China\}} \Delta_h \tau_{jg,1979} \\ & + \sigma^h_{Others} \mathbb{1}_{\{j\neq China\}} \Delta_h \tau_{jg,1979} + \delta_{jh} + \delta_{gh} + u_{jg}, \end{aligned} \tag{3}$$

where $\Delta_h v_{jg,1979}$ is defined as the h-year log difference in import values relative to 1979, i.e., $v_{jg,1979+h} - v_{jg,1979}$, for $h = 1, 2, \ldots, 25$.[13] We follow Boehm et al. (2020) and instrument the h-year change in tariffs relative to 1979 with the tariff change between 1980 and 1979 to account for the autocorrelation properties of this tariff change.[14] The fixed-effect structure is the same as in (2), except that δ_{jg} is eliminated by taking differences of the dependent variable.

[12] For example, in appendix C we document that U.S. imports from Vietnam responded similarly when Vietnam was granted NTR status in 2002. Figure F.2 shows that while the long-run response was similar to China's, Vietnam's adjustment occurred faster. We attribute this to the absence of uncertainty regarding the permanence of this status, and perhaps a greater degree of anticipation that it would be granted.

[13] In Alessandria et al. (2021b), we show that a local projection approach leads to a downward bias in the medium- and long-run trade elasticities when years prior to a tariff change are included. Therefore, we include changes relative to 1979 only.

[14] Figure F.1 of the appendix shows that the tariff changes from the NTR access were permanent and if anything slightly increased over time. This contrasts with the predominantly mean reverting tariff changes over the full sample.

The dashed line in panel (a) of Figure 3 shows the path of adjustment to a tariff change implied by our local projections estimation. While the short- and long-run elasticities obtained by the local projections are very close to those of the ECM approach, the non-parametric estimation of the year-by-year adjustment displays a much slower transition. As panel (b) shows, local projections imply a more gradual transition of U.S. imports from China than the ECM. Our local projections estimates imply that it takes 10 years to complete 50 percent of the total adjustment and 20 years to complete 90 percent. The final adjustment (although it is statistically insignificant) occurs after 22 years, precisely when China joined the WTO and gained PNTR status. Appendix A shows that these results, too, are robust to a wide range of alternative specifications.

Overall, these results indicate that the adjustment of U.S. imports from China to the 1980 reform was large and prolonged, consistent with previous studies of other major trade liberalizations. An important caveat is that our estimates of this adjustment path are likely to be confounded by the effects of uncertainty about whether this reform would be reversed. Our specifications (2) and (3) include only changes in current applied tariffs as regressors, but, as discussed in section 2.2, there could also have been changes in expectations about future tariffs. If the 1980 reform was initially viewed as unlikely to be permanent, as we find in our quantitative analysis, the initial trade response to this reform was likely smaller than it would have been in the absence of this uncertainty, and the adjustment process longer.

2.4 Effects of the risk of losing NTR status

Our second empirical approach draws from the trade-policy uncertainty (TPU) literature, particularly the seminal studies of Pierce and Schott (2016) and Handley and Limão (2017). These studies document that U.S. imports from China grew rapidly during the decade around China's WTO accession even though U.S. tariffs on Chinese goods did not change relative to tariffs on other WTO members. When China joined the WTO, however, the United States lost the ability to revoke China's NTR status and imports of goods that had faced the largest tariff risk consequently grew fastest. This is taken as evidence that the risk of future tariff increases depressed trade and that eliminating this risk stimulated trade growth. Our contribution lies in showing how the effect of tariff risk changed over the history of U.S.-China trade, going all the way back to—and even before—China was first granted NTR status in 1980.

The unifying theme in the TPU literature is a difference-in-difference empirical strategy that compares goods or industries that were more exposed to uncertainty about future policy to goods or industries that were less exposed. In the U.S.-China context, exposure to policy uncertainty is typically measured by the NNTR gap, which is defined as the difference between NNTR and MFN tariffs. Our estimating equation is

$$v_{jgt} = \sum_{t=1974}^{2007} \beta_t \mathbb{1}_{\{t=t' \wedge j=China\}} GAP_g + \delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}. \tag{4}$$

We follow Pierce and Schott (2016) and use the NNTR gap in 1999 as a time-invariant measure of the NNTR gap: $GAP_g = \tau_g^{NNTR} - \tau_{g,1999}^{MFN}$.[15] We include the same set of fixed effects as in section 2.3. Our coefficient of interest, β_t, measures how much the exposure to TPU, as measured by the NNTR gap, lowered U.S. imports from China, on average, each year relative to imports of the same good in 2008 and relative to imports of the same good from other NTR countries. In what follows, we refer to this coefficient as the *elasticity of trade to the NNTR gap*, or the *NNTR-gap elasticity*.

Figure 4 shows the results of this analysis. Between 1974 and 1979, the NNTR-gap elasticity was relatively stable around -10, indicating that imports of high-gap goods were significantly depressed relative to imports of low-gap goods before the NTR liberalization in 1980. Contrary to the conventional interpretation of the NNTR gap in the TPU literature, this effect cannot be attributed to the risk of losing NTR status because this status had not yet been attained. The NNTR-gap elasticity during this period was high simply because tariffs fell less on high-gap goods than on low-gap goods in 1971, so imports of the former initially grew less than imports of the latter. To formalize this point, Figure 5 shows that the NNTR gap in 1999 is highly correlated with the changes in

[15] Variation over time in the NNTR gap is driven by changes in MFN tariff rates; NNTR tariffs, which were set in 1930 by the Smoot-Hawley Act, are constant. Since we exclude countries other than China that did not have NTR status throughout our entire sample, the good-time fixed effects in our specification absorb the effects of changes in MFN tariffs. In the online appendix, we show that our results are robust to using the mean gap over the sample period as well as a time-varying NNTR gap.

applied tariffs on Chinese goods that occurred between 1979 and 1981.[16] Note that this also explains why the NNTR-gap elasticity prior to 1980 is very similar to the long-run trade elasticity we estimated in section 2.3 above.

When China gained NTR status in 1980, the NNTR-gap elasticity rose sharply and then leveled off for several years, and did not begin to grow steadily until 1986. Our interpretation of this finding is that gaining NTR status initially caused high-gap imports to increase because tariffs on these goods fell relative to tariffs on other goods, but this reform was initially perceived as likely to be reversed. Then, in the late 1980s, the credibility of China's NTR status began to rise, leading to a sustained boom in high-gap imports despite the fact that there was no material change in tariffs. This interpretation is confirmed by our quantitative analysis in section 3, and is consistent with the finding of Bianconi et al. (2021) that stock returns of U.S. firms in industries with high NNTR gaps fell during this period. The NNTR-gap elasticity again leveled off from 1992 until 1998, before rising again in the lead-up to WTO accession and PNTR status. This second slowdown in high-gap import growth began shortly after the U.S. Congress started voting annually on China's NTR status in the wake of the 1989 Tiananmen Square incident, which is widely cited as a key event in the TPU literature. Our results indicate, though, that the effect of this event on the NNTR-gap elasticity was relatively small in comparison to the changes that occurred during the 1980s. This suggests that the changes in the risk of losing NTR status played a more important role in the growth of U.S.-China trade immediately after China was granted NTR status in 1980 than in the period surrounding WTO accession. In fact, our estimates show that the change in the NNTR-gap elasticity between the late 1990s and early 2000s was statistically indistinguishable from zero.

These results, too, are robust to a wide range of alternative specifications. Perhaps most importantly, we have estimated the annual NNTR-gap elasticities using more aggregated global data on bilateral imports and exports (not just U.S. imports) to consider the role of good-specific Chinese supply factors. This allows us to control for any spurious correlation between the NNTR gap and differences across goods in export licenses,

[16] A linear regression yields a coefficient of 0.82 and an R-squared of 0.88. The correlation would be perfect but for the modest reductions in MFN tariffs that have occurred since 1981.

state-owned enterprises, import quotas, spillovers from infrastructure development, etc. Appendix B presents the results of these sensitivity analyses.

Just as our estimates of the pace of adjustment to the 1980 NTR status grant in section 2.3 are likely to be confounded by the effects of the risk that this grant could be reversed, our estimates here of the effect of this risk are likely to be confounded by the slow adjustment process. As discussed above, the large NNTR-gap elasticity during the 1970s indicates that the NNTR gap captures the effects of the original tariff reduction that occurred in 1980 as well as the risk that this reduction would be reversed later on. The extent to which the NNTR gap captures the lagged effect of the NTR status grant is likely to diminish as one moves forward in time. To illustrate this point, Figure 4 also shows the results from a version of (4) in which lagged import volumes, $v_{gj,t-1}$, are included in the list of regressors as in (2). Conditioning on lagged imports shrinks the NNTR-gap elasticities by more than 50 percent in the 1970s and 1980s and the estimates from this specification remain significantly (in a statistical sense) smaller than the estimates from the baseline specification until the 1990s. This suggests that the lagged effect of the 1980 reform remained an important driver of the NNTR-gap elasticity until more than a decade later, consistent with our findings in section 2.3. Note that the slowdown in the growth of high-gap imports during the 1980s is even more evident when controlling for lagged trade—the NNTR-gap elasticity in this specification actually reverses direction after the initial spike in 1980—which further suggests that uncertainty may have been particularly important early on.

3 Model and calibration

To quantify the roles of gradual adjustment and policy uncertainty to the the growth of U.S. imports from China, we calibrate a structural model to match the empirical evidence documented above. Our model builds on Alessandria et al. (2021a) and Handley and Limão (2017). Heterogeneous firms pay sunk costs to begin exporting and gradually become better at exporting over time, which generates persistence in export participation at the micro level and gradual adjustment to tariff changes at the macro level. There are two trade policy regimes, one with high tariffs and one with low tariffs, and the probability of switching between regimes varies over time, generating a time-varying effect of tariff risk on trade. We identify the parameters that govern these features using indirect inference (Gourieroux et al., 1993). Specifically, we choose values for these parameters so that running the reduced-form regressions in sections 2.3 and 2.4 on simulated data from the model yields the same coefficients as the actual data.

3.1 Model

The model consists of G goods that correspond to the 5-digit SITC goods in our 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,

$$y_t = z_t \ell_t, \tag{5}$$

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,

$$d_{gt}(p, \tau_{gt}) = \left(p\tau_{gt}\right)^{-\theta} D_{gt}, \tag{6}$$

where τ_{gt} is the current U.S. tariff on goods of type g and D_{gt} 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 that the probability a firm retains its current ξ is symmetric: $P(\xi_L|\xi_L) = P(\xi_H|\xi_H) = \rho_\xi$. This specification implies exporters start exporting a small share of sales and, with some luck and repeated investments, grow to have a large export intensity. 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 but are a function of the firm's export history. 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 $f(\xi_L) = f(\xi_H) = f_1$. This model generalizes the sunk cost model of Das et al. (2007) in a way that can capture the exporter lifecycle (Ruhl and Willis, 2017).

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 that 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 probabilities: $\{\omega_{21,t}\}_{t=0}^{\infty}$ and $\{\omega_{12,t}\}_{t=0}^{\infty}$. We consider alternative information structures in section 4 and appendix D.

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,

$$\pi_{gt}(z_t, \xi_t, \tau_{gt}) = \max_{p,\ell} p\, d_{gt}(p, \tau_{gt}) - w_t \ell \tag{7}$$

$$\text{s.t.} \;\; z_t \ell \geq d_{gt}(p, \tau_{gt}) \xi_t. \tag{8}$$

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

$$V^1_{gt}\left(z_t, \xi_t, \tau_{gt}\right) = -f(\xi_t) + \frac{\delta(z_t)}{1+r} \mathop{\mathbb{E}_t}_{z,\xi,\tau_g} V_{g,t+1}\left(z_{t+1}, \xi_{t+1}, \tau_{g,t+1}\right), \tag{9}$$

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

$$V^0_{gt}\left(z_t, \xi_t, \tau_{gt}\right) = \frac{\delta(z_t)}{1+r} \mathop{\mathbb{E}_t}_{z,\tau_g} V_{t+1}\left(z_{t+1}, \infty, \tau_{g,t+1}\right), \tag{10}$$

and the value of the firm is

$$V_{gt}\left(z_t, \xi_t, \tau_{gt}\right) = \pi_{gt}(z_t, \xi_t, \tau_{gt}) + \max\left\{ V^1_{gt}\left(z_t, \xi_t, \tau_{gt}\right), V^0_{gt}\left(z_t, \xi_t, \tau_{gt}\right) \right\}. \tag{11}$$

The break-even exporter, who is indifferent between exporting and not exporting, has productivity $\bar{z}_{gt}(\xi)$ such that

$$V^1_{gt}\left(\bar{z}_{gt}(\xi), \xi, \tau_{gt}\right) = V^0_{gt}\left(\bar{z}_{gt}(\xi), \xi, \tau_{gt}\right), \tag{12}$$

which can be rewritten as

$$f(\xi_t) = \frac{\delta(z)}{1+r} \left\{ \mathop{\mathbb{E}_t}_{z,\xi,\tau_g} \left[V_{t+1}(z_{t+1}, \xi_{t+1}, \tau_{gt+1}) \right] - \mathop{\mathbb{E}_t}_{z,\tau_g} \left[V_{t+1}(z_{t+1}, \infty, \tau_{gt+1}) \right] \right\}, \tag{13}$$

which says that for firms at the margin, the fixed cost of exporting equals the expected present value of the gain in firm value from exporting in the future. Crucially, this latter object depends on the entire expected path of future tariffs, not the current applied tariff rate.

Aggregation. The decision rules determine how the distribution of productivity and variable trade costs across firms, $\varphi_{gt}(z,\xi)$, evolves over time. The law of motion for this distribution is given by

$$\varphi_{g,t+1}(\mathcal{Z},\infty) = \sum_{\xi}\left[\int_0^{\bar{z}_{gt}(\xi)} h(\mathcal{Z},z)\varphi_{gt}(z,\xi)dz + \int_{\bar{z}_{gt}(\xi)}^{\infty} \bar{h}(\mathcal{Z})\varphi_{gt}(z,\xi)dz\right], \tag{14}$$

$$\begin{aligned}\varphi_{g,t+1}(\mathcal{Z},\xi_H) &= \int_{\bar{z}_{gt}(\infty)}^{\infty} h(\mathcal{Z},z)\varphi_{gt}(z,\infty)dz + \rho_\xi \int_{\bar{z}_{gt}(\xi_H)}^{\infty} h(\mathcal{Z},z)\varphi_{gt}(z,\xi_H)dz \\ &\quad + (1-\rho_\xi)\int_{\bar{z}_{gt}(\xi_L)}^{\infty} h(\mathcal{Z},z)\varphi_{gt}(z,\xi_L)dz,\end{aligned} \tag{15}$$

$$\varphi_{g,t+1}(\mathcal{Z},\xi_L) = (1-\rho_\xi)\int_{\bar{z}_{gt}(\xi_H)}^{\infty} h(\mathcal{Z},z)\varphi_{gt}(z,\xi_H)dz + \rho_\xi \int_{\bar{z}_{gt}(\xi_L)}^{\infty} h(\mathcal{Z},z)\varphi_{gt}(z,\xi_L)dz. \tag{16}$$

where $\mathcal{Z}$ is a typical subset of $\mathbb{R}_{++}$, $h(\mathcal{Z},z)$ is the probability of surviving and drawing a new productivity in $\mathcal{Z}$ conditional on today's productivity z, and $\bar{h}(\mathcal{Z})$ is the probability of dying and being replaced by a new firm with productivity in $\mathcal{Z}$. Although the decision rules, $\bar{z}_{gt}(\xi)$, respond immediately to trade policy changes, the export participation rate adjust gradually as new firms draw good enough productivity shocks to begin exporting. The measure of low-variable-cost exporters adjusts even more slowly, because variable trade costs are persistent and new exporters start with high costs. Consequently, aggregate trade volumes,

$$EX_{gt} = \sum_{\xi\in\{\xi_L,\xi_H\}} \int_z p\left(z,\xi,\tau_{gt}\right) y\left(z,\xi,\tau_{gt}\right) \varphi_{gt}\left(z,\xi\right) dz, \tag{17}$$

respond slowly to policy changes. This is precisely what makes our model well-suited to measuring the roles of gradual adjustment and policy uncertainty in accounting for the growth of U.S. trade with China.

3.2 Calibration

We initialize the model in 1970 to have zero exporters (that is, all firms have $\xi = \infty$) and feed in the realized path of tariffs (NNTR rates during 1971–1979 and MFN rates from 1980 onward). We then calibrate the model's parameters so that (i) its steady state matches salient facts about export participation from the literature; and (ii) its transition dynamics match the empirical evidence on the growth of U.S.-China trade documented in sections 2.3 and 2.4. The calibrated parameter values are summarized in Table 2.

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. 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

$$\log a_{t+1} = \rho_z \ln a_t + \varepsilon, \qquad \varepsilon \overset{iid}{\sim} N(0, \sigma_z^2), \tag{18}$$

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 $1 - \delta\left(a\right) = \max\left[0, \min\left(e^{-\delta_0 a} + \delta_1, 1\right)\right]$. The parameters that govern the firm's productivity and survival processes, ρ_z, σ_z, δ_0, and δ_1, are taken from Alessandria et al. (2021a).

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 (2) on simulated data from the model reproduces the short- and long-run trade elasticities for China of -2.29 and -7.96, respectively, as reported in Table E.1.[17] The identification of these parameters works as follows.

[17] Alternatively, we could target the short- and long-run trade elasticities from our local projections spec-

Parameter θ governs the response of trade to a change in tariffs holding fixed the measures of low- and high-capacity exporters (firms with $\xi = \xi_H$ or $\xi = \xi_L$). These measures do not change immediately when tariffs change, so θ is identified by the short-run trade elasticity. ρ_ξ governs the measure of high-capacity exporters (firms with $\xi = \xi_L$) in a stationary equilibrium, so it is identified by the long-run elasticity. We find $\theta = 3.55$ and $\rho_\xi = 0.87$.[18]

Parameters calibrated to annual NNTR gap elasticities. The probabilities of switching between tariff regimes, $\omega_{12,t}$ and $\omega_{21,t}$, are chosen so that estimating equation (4) on simulated data from the model reproduces the annual NNTR-gap elasticities shown in Figure 4. 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 1974–1979. 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.[19] 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. 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.

In panel (a) of Figure 6, we plot the annual NNTR gap elasticities from the model and

ification (3). Since these elasticities are very similar to the ECM estimates, the implied parameter values would also be very similar.

[18] Note that 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 linearly from 1979 through 1981, as shown in Figure 4.

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

the data. The elasticities from the data are from Figure 4. 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 Results

Having laid out the model and described its calibration, we turn now to the results of our quantitative analysis. First, we discuss how our estimates of trade-policy expectations have evolved since 1980. Second, we study the effects of alternative expectations about future trade policy on U.S. imports from China. We consider a model with no uncertainty about trade policy and we explore a model with a limited number of changes in trade policy expectations around pivotal dates. Third, we study the importance of slow adjustment to the 1980 reform by analyzing a simpler version of our model without an exporter life cycle. Last, we compare our model-generated probabilities of losing MFN status with those from a model-free Bayesian learning approach.

4.1 Estimates of trade-policy expectations

Our estimates show that China gaining NTR status was perceived as unlikely after the embargo ended in 1971, and that this status initially was expected to be quickly revoked after it was granted in 1980. The main change in expectations about the persistence of NTR status took place from 1986 through 1993. Since then, the likelihood of keeping MFN tariffs has been very high and that persistence was not altered materially by China joining the WTO, especially compared with the changes in expectations that took place in the late 1980s.

Panel (b) of Figure 6 plots the main results of our calibration exercise: the estimated probabilities of switching between policy regimes. The probability of switching from the NNTR to the MFN regime is about 0.25. The probability that the economy switches back to the NNTR regime is initially 0.72 in 1980, rises to 0.81 in 1981, and then begins to fall throughout the mid 1980s and early 1990s, even following the Tiananmen Square incident in 1989 and the subsequent introduction of Congressional voting.[20] 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 that there is no discrete drop or accelerated decline in this probability when China gains PNTR status in 2000 (in fact, the switching probability rises slightly, from 5.3 percent to 6.3 percent, during 2000–2002). Thus, 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 unlikely even before China gained PNTR status. This finding reflects the observation in section 2.4 that this change in status had little impact on the annual NNTR-gap elasticity.

Another way to interpret these results is to compare the realized path of tariffs to the expected present value of tariffs that Chinese exporters face at each point in time. Panel (c) of Figure 6 plots the mean expected present value of tariffs across goods in the model,

$$\tau_t^{PV} = \frac{1}{G}\left\{\sum_{g=1}^{G}(1-\beta)\left(\sum_{s=t}^{\infty}\beta^{s-t}\mathbb{E}_t[\tau_{gs}]\right)\right\}, \tag{19}$$

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 MFN 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 a small discrete drop in 1980, and the expected present value remains above the applied MFN rate even after China joins the WTO in 2001.

The dynamics of the expected present value of tariffs helps us understand the gradual adjustment of export volumes to the abrupt decrease in the current tariff. Recall that the intensive margin of trade, the exports per exporter, is mostly determined by the current

[20] Our finding of a substantial increase in the probability of MFN renewal from 1986 to 1993 lines up well with several key policy reforms in China. In July 1986, China applied to join the GATT and in March 1987 a working group was formed to examine China's application and negotiate terms of accession. This application followed substantial economic reforms towards a more market-oriented economy in agriculture and industry.

tariff and distribution of variable trade costs, while the extensive margin, through entry and exit decisions, is determined by the path of future tariffs. The large gap between the expected present value of tariffs and current tariffs suppressed the participation of Chinese firms in the export market. As the expected present value of tariffs fell, export participation increased and aggregate trade volumes grew.

4.2 The effects of policy uncertainty on trade

In addition to allowing us to recover the path of expectations about U.S.-China trade policy, our model allows us to measure how changes in these expectations affect trade flows. To do so, we compare our benchmark model with a counterfactual in which firms always believe the current trade-policy regime is permanent (this is equivalent to letting the persistence parameters, ω_{11} and ω_{22}, converge to one). In this no-TPU counterfactual, firms believe the NNTR regime will last forever during 1971–1979. In 1980, firms are surprised by the shift to the MFN regime but believe that the MFN regime will last forever.[21] Because the realized path of tariffs in this model is the same as in the benchmark, differences in trade growth are due solely to differences in the expected path of tariffs.[22]

The no-TPU counterfactual allows us to assess the contribution of trade-policy uncertainty to the NNTR-gap elasticity. Panel (a) of Figure 6 plots the elasticities obtained by estimating (4) on simulated data from the counterfactual against our benchmark elasticities. During the 1970s, the elasticities in the counterfactual are lower than in the benchmark. As discussed in section 3.2, the possibility of gaining MFN status boosts trade in high-gap goods during this period. After 1980, the elasticity in the counterfactual rises faster than in the benchmark, as more firms in high-gap industries enter in response to the perfectly credible reform. By the late 1990s, the gap elasticity in the counterfactual essentially converges to zero. The difference between the counterfactual elasticity and the actual elasticity measures the contribution of policy uncertainty to the latter at each point in time. This contribution peaks at 78 percent during the late 1980s, falls to 60 percent dur-

[21] We have also studied an alternative no-TPU counterfactual in which the 1980 reform is anticipated, rather than unanticipated. This leads to very similar results post-1980, except for a rapid rise in the NNTR-gap elasticity during the late 1970s, as export participation in high-gap industries begins to rise ahead of the 1980 reform.

[22] We do not seek to separately identify the role of expected tariffs from uncertainty about tariffs.

ing the early 1990s, and rises again to 70 percent during the late 1990s. It is interesting to note that trade policy uncertainty accounts for a similar portion of the NNTR-gap elasticity during the periods when NTR reversal was most likely (the late 1980s) and least likely (late 1990s). This is, of course, because the overall elasticity was much smaller during the latter period and the adjustment to the 1980 reform had mostly been completed.

The no-TPU counterfactual also allows us to assess the impact of trade-policy uncertainty on aggregate trade. Panel (d) of Figure 6 shows aggregate exports in the counterfactual and aggregate exports in the benchmark. Again, the vertical distance between the two lines measures the effect of trade-policy uncertainty. The results largely mirror those shown in panel (a) with one notable exception: there is still a material difference in aggregate trade between the counterfactual and the benchmark after China joins the WTO in 2001. This is because the probability of losing NTR status is positive (albeit small) even in the long run. The long-run expected present value of tariffs in the benchmark model is higher than the applied tariff rate (panel c), and this permanently reduces the number of exporters in high-gap industries.

4.3 The role of time-varying policy uncertainty

Our quantitative strategy leverages year-by-year variation in the elasticity of trade to the NNTR gap to identify changes in expectations about U.S. trade policy towards China. Our findings suggest that the growth in U.S. imports from China during the late 1990s and early 2000s was not driven by a large decline in policy uncertainty that occurred when China gained PNTR status in 2001, but rather a slow adjustment to earlier reductions in the likelihood of reverting to NNTR that started in the mid 1980s. Here, we ask: What would happen if we assumed there were only one or two changes in expectations at key geopolitical moments?

We answer this question by first studying a version of our model in which the probability of losing NTR status is constant between 1980 and 2000, and then goes to zero in 2001.[23] We calibrate the probability of losing NTR status during 1980–2000 to match

[23] We assume that firms know in advance that the probability will go to zero in 2001. The results are similar when we treat this change as a surprise, although there is a much sharper spike in the NNTR gap elasticity after 2001 that is at odds with the data shown in Figure 4.

the single pre-PNTR elasticity of trade to the NNTR gap estimated by Pierce and Schott (2016) over the period 1992–2007, i.e., the coefficient β from the regression

$$v_{jgt} = \beta \mathbb{1}_{\{t<2000 \wedge j=China\}} X_g + \sigma \tau_{jgt} + \delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}. \tag{20}$$

Appendix B provides additional details about this regression. Panel (a) of Figure 7 shows that this version of the model (*Const. TPU from 1980*) does fairly well in matching the evolution of the NNTR-gap elasticity during the 1990s and 2000s. Trade grows too quickly, however, in high-gap industries relative to low-gap industries in the 1980s. Panel (b) shows that the probability of losing NNTR status in this version of the model is almost 25 percent, much higher than in the benchmark. This analysis shows that rising credibility of U.S. trade policy towards China during the mid-to-late 1980s was an important factor in explaining the growth of U.S. imports from China, and that ignoring this trend leads one to overstate the degree to which uncertainty fell after China gained PNTR status.

We also consider a version of this analysis in which there is no risk of losing NTR status prior to 1990 but a constant risk in the 1990s. This exercise is intended to capture the idea that non-renewal did not become a serious concern until Congress started to vote annually on China's NTR status in the wake of the Tiannanmen Square incident in 1989, which is a common assumption in the trade-policy uncertainty literature. For example, Bianconi et al. (2021) write that "annual renewals by Congress... were essentially automatic until the Tiananmen Square Massacre in 1989. Starting in 1990, NTR renewal in Congress became more politically contentious..." (see also Section I.A in Pierce and Schott, 2016). Panel (a) of Figure 7 shows that this model (*Const. TPU from 1990*) also does a good job of matching the trajectory of the NNTR-gap elasticity from about 1995 onward. However, its behavior during the 1980s and early 1990s is sharply at odds with the data. Here, the elasticity shrinks even faster during the 1980s—it follows the no-TPU counterfactual's trajectory exactly—before growing again during the 1990s when the risk of losing NTR status is realized. The non-monotonic behavior of the NNTR-gap elasticity in this model is clearly at odds with the data. China's export growth cannot be understood without time-varying uncertainty about the persistence of China's NTR status, particularly in the early years after this status was granted.

4.4 The role of slow adjustment

Our estimates of trade-policy transition probabilities depend on the speed of trade adjustment that our model of exporter dynamics generates. When we estimate a simpler conventional model with sunk export costs, we find much larger estimates of the non-renewal probability and a much smaller likelihood of transitioning from MFN to NNTR.[24] This is because the simpler model leads to faster transitions than in our benchmark model.

In panel (a) of Figure 8, we plot the elasticity of trade to the NNTR gap in the baseline model, the baseline model without TPU, and the fast adjustment model without TPU. The *fast adjustment model* replaces the stochastic variable trade cost with a constant variable trade cost. Thus, a new exporter immediately exports at its full scale and aggregate trade responds quickly to a change in policy. The elasticities in the fast adjustment model without TPU converge to zero faster than in the baseline model without TPU. As panel (b) of Figure 8 shows, this implies that higher probabilities of switching back to the NNTR regime are required to match this trajectory—in fact, the simple dynamic model requires this probability to be one in 1980–1981. Trade adjusts too quickly in the fast adjustment model, and as a result, this model places too much weight on expectations and too little on slow adjustments in accounting for the dynamics of U.S. imports from China. Modeling the slow adjustment we observe in the data is crucial for accurately measuring expectations about trade policy.

4.5 Bayesian learning

The path of trade policy expectations that is consistent with the growth of Chinese exports to the United States reflects complex geopolitical events that are beyond the scope of our analysis. In our model, firms do not update their beliefs about the probabilities of switching between regimes; they simply take the probabilities that are "announced" by the modeler. Here, we ask how the probabilities we obtain from our calibration exercise compare with the posterior beliefs that a Bayesian agent would form after observing the

[24] This model is quite close to that analyzed by Handley and Limão (2017).

economy remain in the MFN regime year after year from 1980 onward. Surprisingly, this alternative approach to forming beliefs about future trade policy yields a path of trade policy expectations that falls with time at a rate that is roughly consistent with our estimates.

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:

$$p^{prior}(\omega_{12}|a,b) = \frac{\Gamma(a+b)}{\Gamma(a)+\Gamma(b)}\omega_{12}^{a-1}(1-\omega_{12})^{b-1}. \tag{21}$$

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. This setup allows us to determine whether agents in our model "learn" faster or slower than a Bayesian agent would. 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 9 plots the density functions of each of the prior beliefs that we consider.

Panel (b) 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 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 2.4. After 1985, however, this pattern is reversed, and by the late 1990s the model-implied probability of losing MFN status is lower than all of the posteriors.

5 Conclusions

We study, empirically and quantitatively, the growth of China's exports to the United States since the long-standing embargo on Chinese goods was lifted in 1971. We find that the dynamics of this integration are consistent with substantial uncertainty about the future path of tariffs, particularly during its initial phase. During the late 1970s, the likelihood of gaining access to U.S. markets at NTR rates was perceived to be low, and once this access was granted in 1980 it was perceived as likely to be revoked. During the mid 1980s, the probability of losing NTR access fell dramatically and it remained low through the late 1990s in the lead-up to China attaining permanent normal trade relations in 2000. This suggests that much of the growth in trade in products with high NNTR gaps in the 1990s and 2000s was a delayed effect of earlier liberalizations—and earlier increases in their perceived credibility.

Our approach to estimating the dynamics of trade-policy expectations leverages unique aspects of U.S. policy toward China in which the potential change in trade policy that could occur in the future is known and heterogeneous across products, while the likelihood of this change is unknown and 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, one could discipline the process for these possible trade-policy outcomes.

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

Finally, that trade simultaneously depends on past, present, and future changes in trade policy suggests we need to rethink our approach to measuring the response of trade to these changes. Alessandria et al. (2021b) build on our findings here to show how to measure the response of trade to unanticipated versus anticipated changes in trade policy as well as policy changes that feature several forms of uncertainty.

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Table 1: Summary statistics for NNTR and NTR tariff schedules

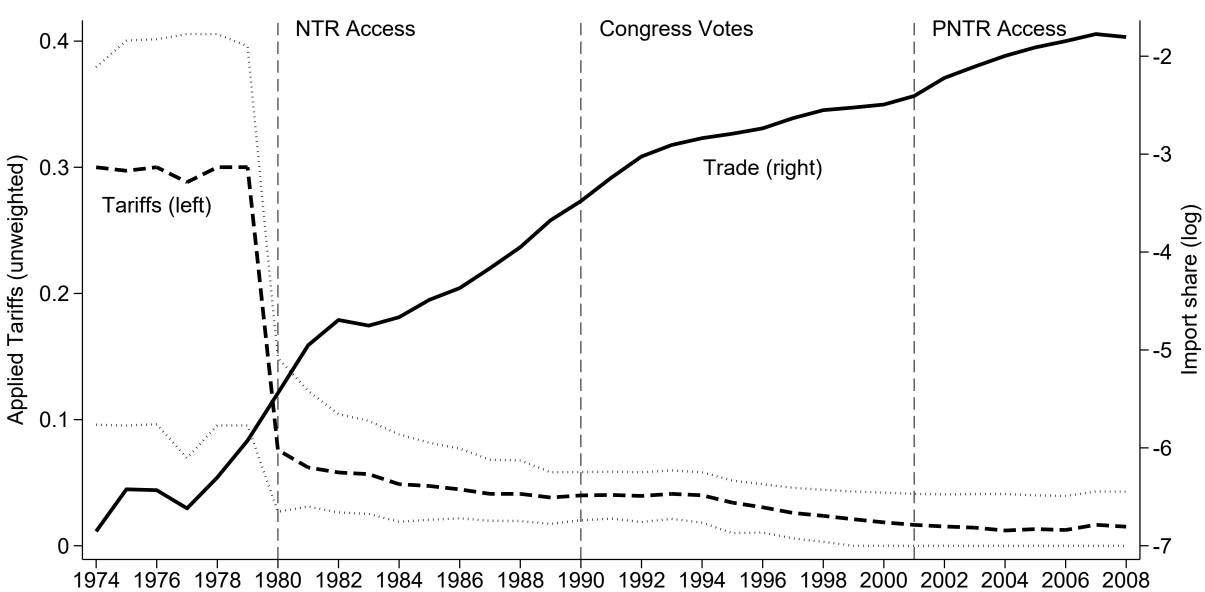
	NNTR Rate		U.S. Export Share		Applied Duties			
					1979		2001	
SITC 1-Digit	Mean	Std.	1979	2001	China	NTR	China	NTR
0 Food and live animals (food)	17	15	10	1	12	5	6	2
1 Beverages and tobacco	48	41	0	0	31	11	4	3
2 Crude materials, inedible, except fuels	11	15	13	1	6	1	1	0
3 Mineral fuels, lubricants 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: All values are in percent. *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 7). When computing the mean applied duties of NTR countries, only SITC codes with nonzero U.S. imports from China are considered.

Table 2: Calibrated parameter values

Parameter	Meaning	Value	Source/target
(a) Assigned			
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	Corr.(survival,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 ECM trade elasticities			
θ	Demand elasticity	3.55	ECM estimate of SR trade elasticity = -2.29
ρ_ξ	Prob. of keeping iceberg cost	0.87	ECM estimate of LR trade elasticity = -7.96
(d) Calibrated to annual NNTR-gap elasticity			
$\omega_{21,t}$	Prob. NNTR to MFN	0.25	Avg. NNTR-gap elasticity during 1974–1979
$\omega_{12,t}$	Prob. MFN to NNTR	Varies	NNTR-gap elasticity during 1980–2008

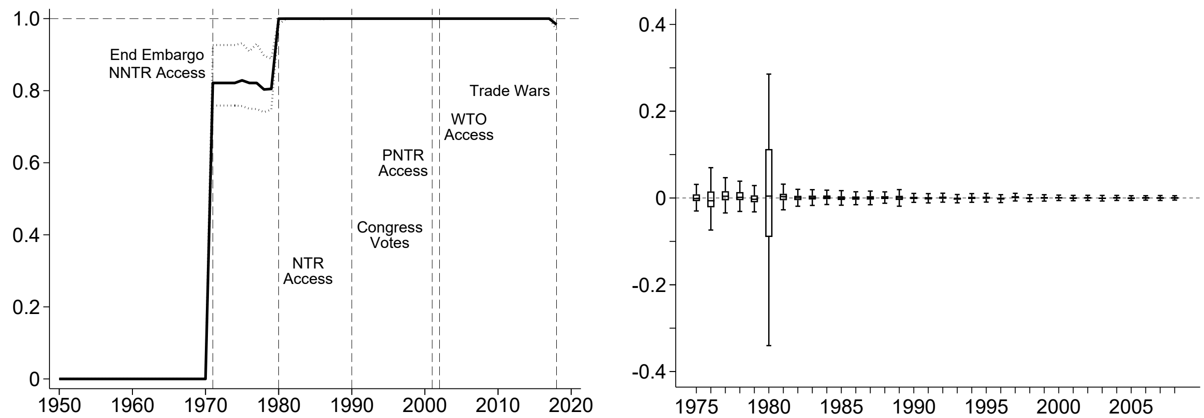
Figure 1: China's aggregate import share and duties



Notes: 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 2: U.S. trade policy toward China vs. other NTR countries

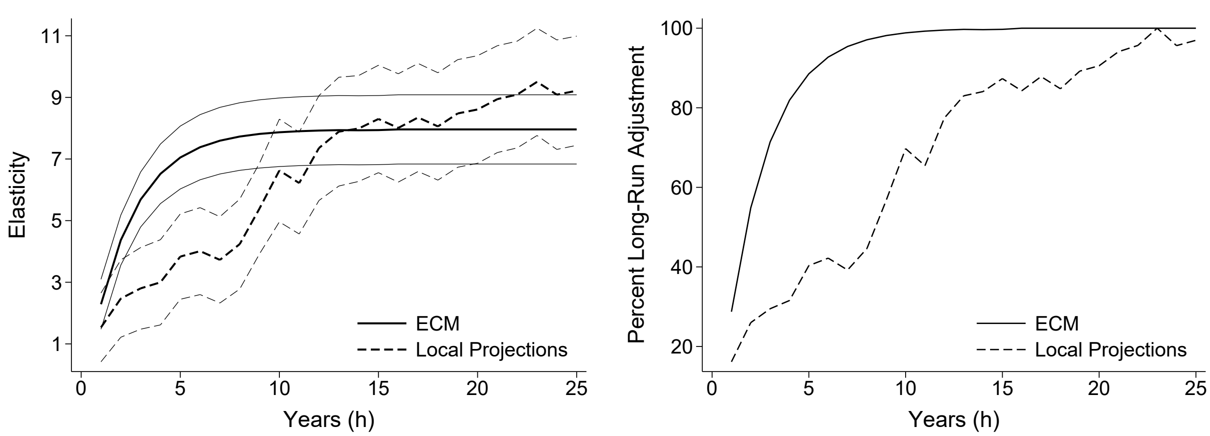
(a) Inverse tariff (b) Residual variation of tariff changes



Notes: Panel (a) plots China's inverse tariff defined in (1). The solid line is the median of all products and the dashed lines are the 25th and 75th percentiles. Panel (b) plots the distribution of the annual residual variation in China's tariff changes by considering the residual from regressing $\Delta_1 \tau_{jgt}$ on δ_{jt}, δ_{jg} and δ_{gt}. The box spans the the lower to the upper quartiles and the line in the box is the median; the upper and lower whisker are the respective adjacent values, i.e. within 1.5 times the interquartile range. Outliers are excluded from the plot.

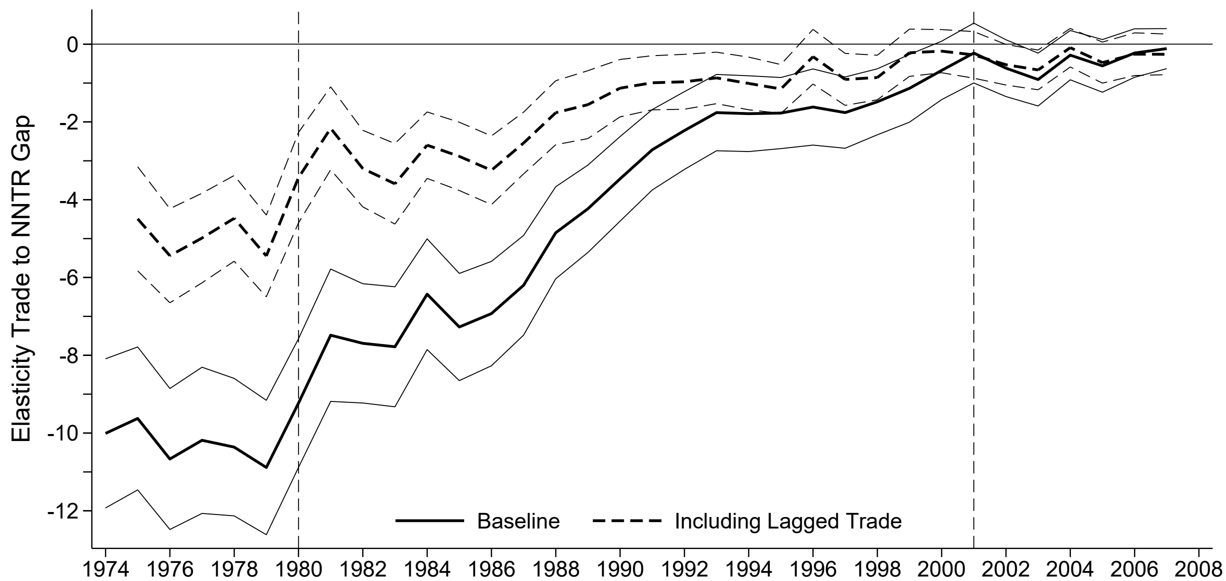
Figure 3: Slow adjustment to tariff changes

(a) Trade Elasticities (b) Speed of Adjustment



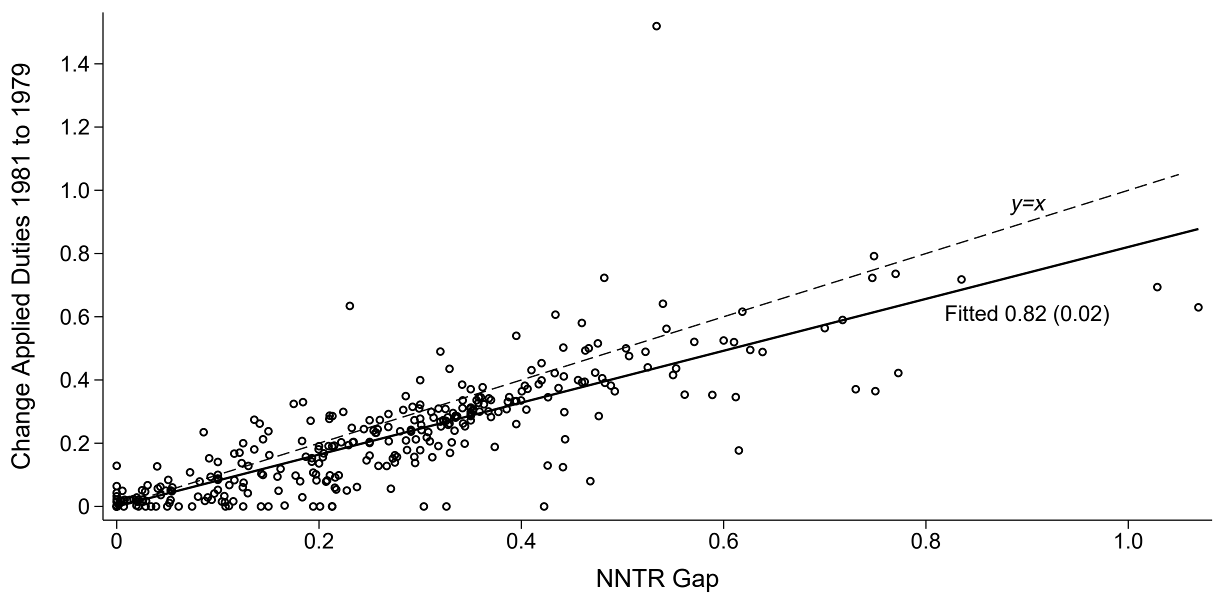
Notes: Panel (a) reports the time varying elasticities to tariff changes on U.S. imports from China. The solid line uses the ECM approach of (2) and the dashed line use the local projection approach of (3). The standard errors that construct the 95 percent confidence intervals are clustered at the gj level. Panel (b) plots the ratio of the elasticity at each period relative to the maximum elasticity throughout all periods.

Figure 4: Elasticity of U.S. imports from China to NNTR gap



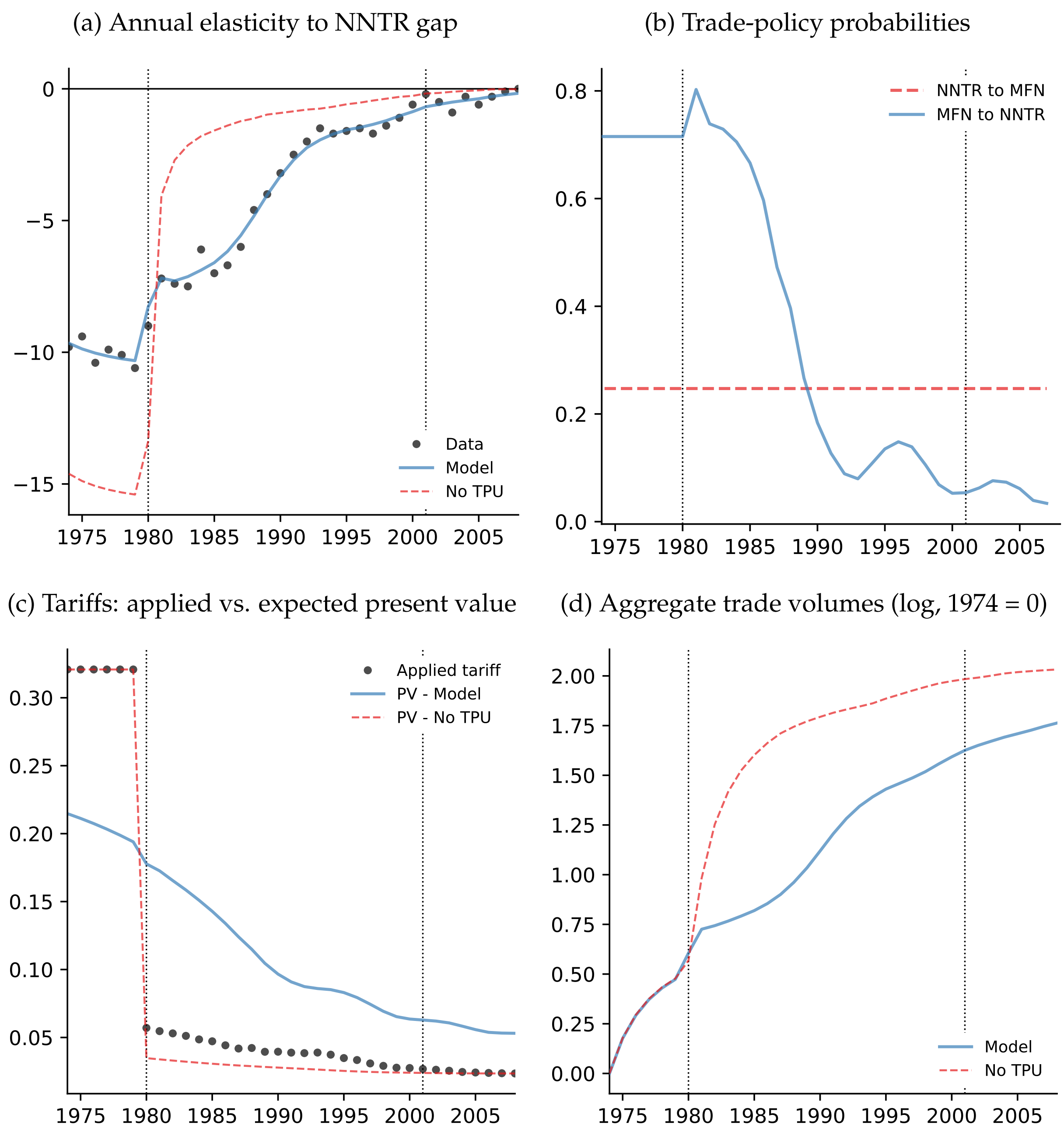
Notes: This Figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (4). The solid line is our baseline estimate. The dashed additionally includes the lagged value of imports $v_{jg,t-1}$ on the right-hand side of (4). The standard errors that construct the 95 percent confidence intervals are clustered at the gj level.

Figure 5: Size of 1980 liberalization vs. NNTR gap



Notes: Each dot is a SITC good (excluding MFA goods). The slope of fitted line (solid line) is the coefficient of regressing the change in applied duties on China's imports between 1979 and 1981 on the NNTR gap (at SITC level). The coefficient is 0.82 and its standard error 0.02. The R-Square of this regression is 0.86.

Figure 6: Trade and policy dynamics in the model with and without policy uncertainty

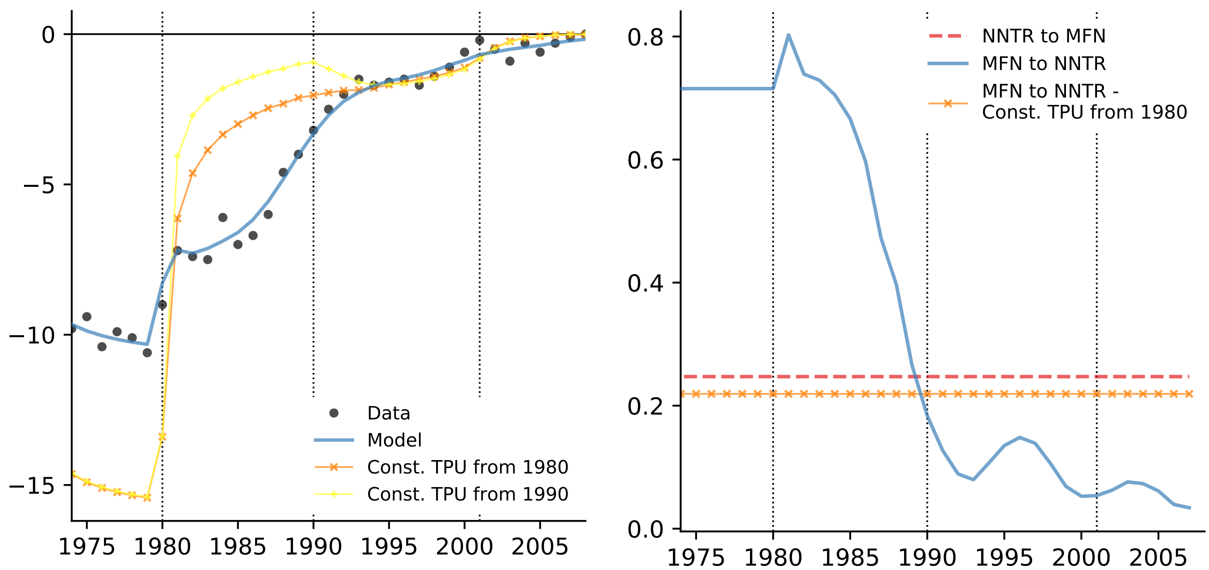


Notes: Panel (a) shows NNTR-gap elasticities estimated using simulated data from the model. Panel (b) shows estimated probabilities of switching policy regimes. Panel (c) shows expected present value of tariffs implied by these probabilities. Panel (d) shows aggregate trade volumes in the simulated data.

Figure 7: Benchmark model vs. constant-TPU models

(a) Annual elasticity to NNTR gap

(b) Trade-policy probabilities

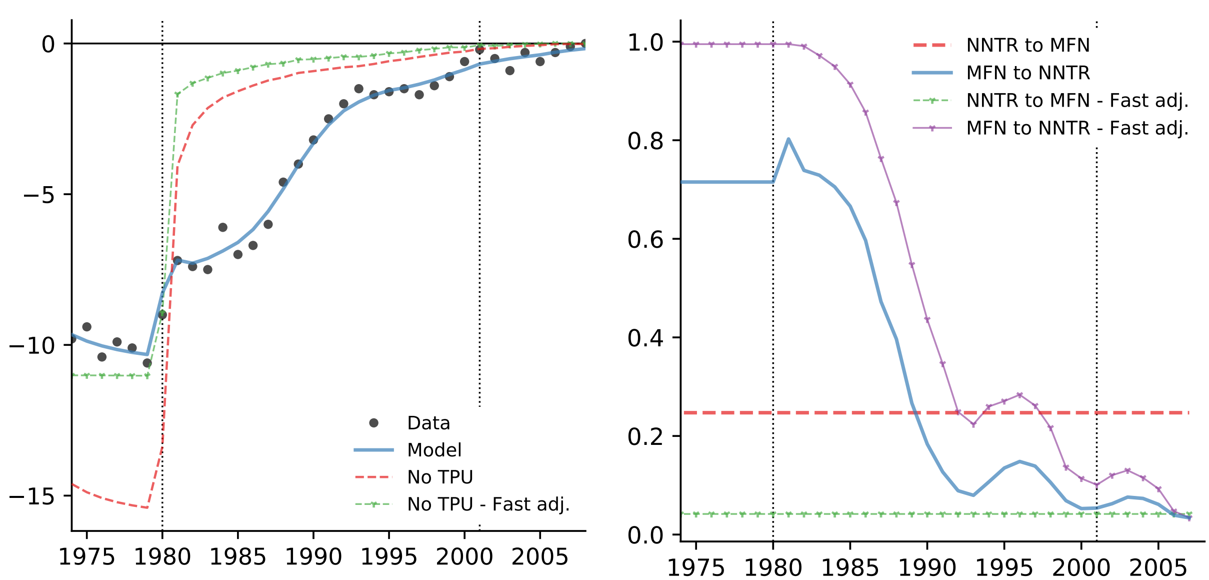


Notes: Panel (a) shows NNTR-gap elasticities estimated using simulated data from the model. Panel (b) shows estimated probabilities of switching policy regimes. In constant-TPU models, probability of switching from the MFN regime to the NNTR regime is constant from either 1980 onward or 1990 onward. Implied probability of losing MFN status is similar in both versions so panel (b) only shows estimate from former.

Figure 8: Benchmark model vs. fast-adjustment model

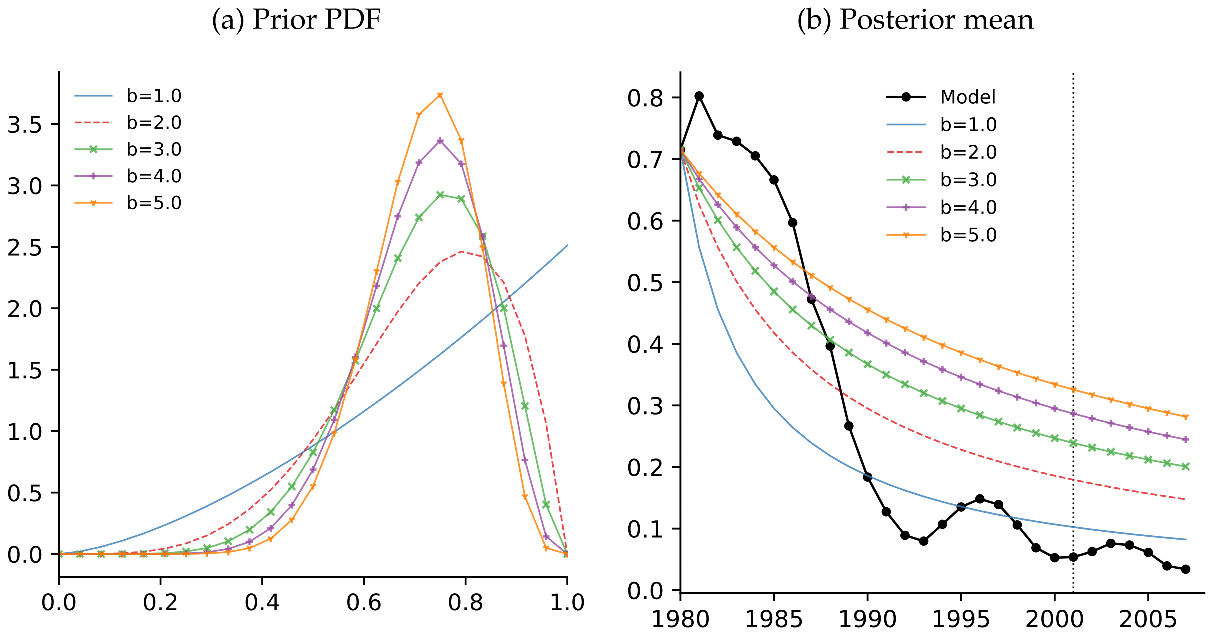
(a) Annual elasticity to NNTR gap

(b) Trade-policy probabilities



Notes: Panel (a) shows NNTR-gap elasticities estimated using simulated data from the model. Panel (b) shows estimated probabilities of switching policy regimes. In fast-adjustment model there are no idiosyncratic shocks to firms' productivities or variable trade costs.

Figure 9: Model-implied expectations versus Bayesian learning



Notes: Panel (a) shows Bayesian prior belief distributions about the probability of losing NTR status. All priors have the same mean as the model-implied estimate for 1980. Panel (b) shows mean posteriors from 1980–2008 against the model-estimated probabilities.

Appendix (For Online Publication)

In appendix A, we document the robustness of our results concerning short- and long-run responses of trade to tariff changes and the particularly gradual adjustment of U.S. imports from China. In appendix B, we document the robustness of our results about the annual elasticity of trade to the NNTR gap. In appendix D, we present the results of additional quantitative experiments. Appendices E and F contain the additional tables and figures discussed in the aforementioned sections.

A Robustness: slow adjustment

The results of the ECM specified in (2) and the local projections in (3) presented in section 2.3 are robust to a number of alternative specifications. For visualization purposes we only present the results corresponding to the ECM. Results of the corresponding local projection results are available upon request.

Shipping Costs. Column 2 of Table E.1 reports results from a version of our ECM regression that includes a control for shipping costs (CIF charges).

Sample of Countries and Goods. Columns 3 to 6 of Table E.1 report the ECM results under different specification of the sample of goods and countries. In column 3 we include all goods, including those affected by the MFA. In column 4 we include all countries, not only those granted NTR status and not part of a bilateral FTA with the United States. Column 5 extends the sample to all goods and countries. Finally, column 6 only includes goods that had non-zero U.S. imports from China at some point in the period 1974–1979. Overall we find very similar short- and long-run elasticities, although the inclusion of MFA goods slightly increases the short-run elasticity and the diminishes the long-run elasticity.

Level of Aggregation. Table E.2 presents the results of our ECM regression applied to more disaggregated datasets: 8-digit TS-USA and the HS-8 level. The former is available for the 1974–1988 period while the latter is available for the 1989–2008 period. To facilitate comparison with our baseline results, we also report the results using the 5-digit SITC aggregation over these same time periods. There are two main takeaways. First, in our baseline sample using the SITC good level, the ratio of China's long- to short-run elasticity is smaller when splitting the sample period into two—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, although it slightly increases when using more disaggregate trade flows. This can be seen comparing columns 2 to 3 and 4 to 5. In all cases, the differences between short- and long-run elasticities are statistically insignificant, while the point estimates of their ratio are slightly larger when using TS-USA and HS-8 level data. These findings indicate that (1) using a more disaggregate level of trade flows, if anything, results in slower adjustment; and (2) it is important to use the long sample period to capture the full extent of the gradualness.

B Robustness: NNTR-gap elasticity

The time-varying pattern of the effect of the NNTR gap on China's exports to the United States shown in Figure 4 is robust to a range of alternative approaches. Before we show these results we report the non-time varying average NNTR-gap elasticities estimated under the approach of Pierce and Schott (2016).

Pierce and Schott (2016) replication. In subsection 2.4 we estimated the NNTR-gap elasticities for each year between 1974 and 2007 as the trade effects of the NNTR gap relative to 2008. Instead, in their seminal article, Pierce and Schott (2016) estimate the effect by considering the differences trade in pre- and post-China's WTO accession. Their estimation equation is the following:

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

There are two differences with respect to our approach in (4). First, a single indicator variable for the pre-WTO accession period ($t < 2000$), is used instead of an indicator variable for each year prior to 2008. Second, (B.1) controls for contemporaneous tariff changes by including the current applied tariff rate, τ_{jgt}. Moreover, Pierce and Schott (2016) focus on the sample period between 1992 and 2007. Here we revisit the result of (B.1) when expanding the sample period back to 1974. As mentioned above this requires the use 5-digit SITC goods instead of HS-8 tariff lines as in Pierce and Schott (2016).

Table E.3 reports the results of several different versions of this regression. In Column 1, we use the same 1992–2007 sample as in Pierce and Schott (2016) and estimate $\hat{\beta} = -0.9$.[25] In Column 2, we use the full sample period, 1974 to 2008, and estimate an effect that is almost three times larger. The remaining columns show that this result holds when using measures of the tariff rates China faced during the 1970s instead of the NNTR gap in 1999. Columns 3 and 4 use the statutory NNTR rate as measured by Feenstra et al. (2002) in place of the NNTR gap, while Columns 5 and 6 use the average applied duty during 1974–1979 in our trade data. The estimates of $\hat{\beta}$ in Columns 3 and 5 are similar to the estimate in Column 1 using data from 1992–2007, while the estimates in Columns 4 and 6 are similar to the estimate in Column 2 using data from the full sample period.

Like our annual NNTR-gap elasticity estimates, these results indicate that the growth in U.S. imports from China in high-gap goods after China gained PNTR access in 2000 is likely to be a delayed effect of the 1980 NTR liberalization as well as a consequence of reduced uncertainty about future trade policy.

Anticipation 1979. Figure 1 shows Chinese exports to the United States grew strongly in 1979, but Figure 4, which shows that the elasticity to the NNTR gap fell in 1979, indicates this increase was smaller for high-gap goods than low-gap goods. The weak growth for goods whose tariffs were about to decline the most could reflect anticipatory behavior by importers as emphasized by Khan and Khederlarian (2021). To control for this possibility, we estimate a version of equation (4) including the lead change in applied tariffs ($\Delta\tau_{jg,t+1}$). This control is only relevant in 1979 because changes in applied tariffs in other years were minimal (see Figure 2). Column 2 of Table E.4 shows the results. This inclusion indeed smooths the response of trade flows to the NNTR gap around 1979 (we no longer see a drop in the elasticity), but has no effect in other years.

[25] Pierce and Schott (2016) estimate a value of -0.5. This difference is due to the fact that our level of aggregation is coarser than theirs.

Additional Trade-Cost Controls. The baseline estimating equation (4) departs from Pierce and Schott (2016) by excluding applied tariff rates τ_{jgt} because the NNTR gap is highly correlated with applied tariffs pre-1980 as shown in Figure 5. Columns 3–4 of Table E.4 shows the results. Including applied tariff rates and/or other shipping costs leaves the estimated coefficients virtually unchanged.

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 the life cycle of trade relationships at the good level. Namely, we introduce the following variables in equation (4): 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 E.4 shows the results. The differences between the results from this specification and our baseline results are minor.

Sample of Countries and Goods. Our results are also robust to the restrictions on countries and goods in our baseline sample. The results are virtually unchanged when we include all countries (Column 1 of Table E.5). When we relax the MFA exclusion (Column 2 of Table E.5), the estimated annual elasticities of trade to the NNTR gap fall by 10–20%, but the drop is common across all years, leaving the speed of adjustment and the overall pattern unchanged. Restricting the sample to goods in which U.S. imports from China were non-zero at some point before 1980 also has little effect (Column 3 of Table E.5).[26]

Alternative Measures of the NTR Liberalization. In our baseline specification (4) we estimate the annual effect of the 1999 NNTR gap on U.S. imports from China to facilitate comparison with Pierce and Schott (2016). We have also estimated this equation using the 1999 NNTR rate (not the gap) and the average applied tariff rate during 1974–1979. Using the 1999 NNTR rate yields annual gap elasticities about 10–20 percent smaller in the first years, but the gradual nature of the adjustment is unchanged (Column 4 of Table E.5).

[26]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.

Using the average applied rate during 1974–1979 yields elasticities that are about 5-15 percent smaller in the than the baseline and converge to zero (or statistical insignificance) slightly faster in the last five years before the PNTR liberalization (Column 5 of Table E.5).

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 that includes 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, as we did above in Appendix A. The TS-USA results are shown in Table E.6 and the HS-8 results are shown in Table E.7.

China Supply Effects. Our baseline approach controls for U.S. demand shocks specific to particular goods but not for good-specific Chinese supply shocks, such as export licensing, privatization of state-owned enterprises or infrastructure development. Note that this would be problematic only if these were systematically correlated with the NNTR gap. An additional caveat is that, especially in the early years of our sample period, U.S. imports represent a large share of World trade. In the presence of economies of scale, the opening of U.S. markets to Chinese goods can have spillover effects and increase exports to other destinations, thereby introducing a downward bias in the elasticity estimates. However, we study the importance of potentially confounding supply factors, by using the world Trade dataset from Feenstra et al. (2005) (1974-2000) merged with the BACI Trade database (2001-08).[27] We estimate

$$
\begin{aligned}
v_{ijgt} = & \sum_{t'=1974}^{2007} \beta_t \mathbb{1}\{t=t'\}\mathbb{1}\{i=U.S.\}\mathbb{1}\{j=China\}X_g \\
& + \mathbb{1}\{i=U.S.\}\mathbb{1}\{j=China\}X_g + \delta_{igt} + \delta_{jgt} + u_{ijgt},
\end{aligned}
\tag{B.2}
$$

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

[27]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.

trade 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 results are shown in Figure F.6. The solid line plots the response of U.S. imports of China when considering U.S. imports only and estimating (4)—analogous to the baseline, but in the world trade sample. The dashed line extends the sample to include all trade flows, thus allowing us to include source-good-time fixed effects as in (B.2). The overall pattern is similar to the baseline; in fact, we cannot say the differences between the results from this specification and our baseline are statistically significant.

C U.S.-Vietnam Trade and Trade Policy

The recent history of "de facto" trade policy of the U.S. towards Vietnam is very similar to China's. In February of 1994 the United States lifted its 30-year long embargo on imports from Vietnam and allowed Vietnamese exporters to access the U.S. markets at the NNTR tariff rates. In fact, Vietnam was the last country to effectively export at these rates when it gained NTR status in December of 2001. This status was granted permanently after Vietnam became a member of the WTO in January of 2007. Here we apply the same approach of section 2 to study the dynamic adjustment to the 2002 NTR access and the annual path of NNTR-gap elasticity.

Figure F.2 shows the results of applying the two approaches of section 2.3 to estimate U.S.-Vietnam's trade adjustment to the 2002 NTR access. The response is very similar to China's. While in the case of Vietnam the short-run elasticity is slightly larger, its long-run response is almost the same at around 8.[28] The most important difference lies in the speed of adjustment. In the case of Vietnam's NTR access the adjustment is completed by five years. We view this as evidence of the lack of uncertainty regarding the permanence of Vietnam's NTR status.

Figure F.3 shows the path of the NNTR-gap elasticities when using the same approach

[28] One potential reason that Vietnam's short-run elasticity is relatively high is that the NTR access was largely anticipated. In Alessandria et al. (2021b) we show that anticipation can lead to large upward biases in the short-run elasticity. This explanation is also consistent with the advent of the liberalization being viewed as likely.

as in (4). As in the case of China, prior to its NTR acces, the gap elasticity does not capture the risk of reverting to NNTR rates, since MFN rates had not yet been attained. Instead, the elasticity captures the import growth of goods with low NNTR rates relative to high NNTR rates after the lifting of the embargo in 1994. This elasticity is quite similar China's pre-1980 NNTR gap elasticity. However, Vietnam's adjustment throughout the period of conditional NTR status—between 2002 and 2007—is different from China's: The NNTR-gap elasticity continues to rise throughout the entire period. This suggests that Vietnam's NTR status lacked the distinct changes in uncertainty that characterized China's trade growth in the aftermath of the NTR access. This is further corroborated in the annual path of the NNTR-gap when lagged trade is included on the right hand side of (4) (dashed line in Figure F.3). The annual effect of the NNTR gap on Vietnam's imports to the U.S. is almost entirely flat between 2002 and 2007. Moreover, there is no significant change between 2006 and 2007, indicating that the 2007 PNTR access was already incorporated into exporters expectations.

D Robustness: quantitative analysis

Here, we conduct three additional quantitative experiments. In the first, we estimate upper and lower bounds for our trade-policy probabilities by calibrating our model to match the confidence intervals of our gap-elasticity coefficients instead of the point estimates. In the second, we study a version of our model in which changes in the probability of switching trade policy regimes are unanticipated instead of anticipated. In the third, we study a version of our model with time-varying iceberg trade costs that vary across goods instead of trade policy uncertainty.

D.1 Bounds for our estimates of trade-policy probabilities

To estimate a path of expectations about future US trade policy towards China, we have calibrated the probabilities of switching between policy regimes in our model to match our point estimates of the annual elasticity of trade to the NNTR gap. In our empirical analysis, we also reported 95% confidence intervals around these point estimates. Here, we use these intervals to derive upper and lower bounds for our estimated probabilities.

Leaving all other parameters unchanged, we re-calibrate our model's trade-policy probabilities two more times, once to match the lower bound of the confidence interval shown in figure 4, and once more to match the upper bound. The former yields a lower bound for the probability of switching from NNTR to MFN and an upper bound for the probability of switching back from MFN to NNTR, while the latter yields the reverse.[29] The results are shown in figure F.7.

Our bounds for the probability of switching from NNTR to MFN tariffs (shown in light blue in the figure) are about 13p.p. below/above than our main point estimate. Our bounds for the probability of switching back from MFN to NNTR tariffs (shown in light red) are about 14–17p.p. below/above our estimate during the 1970s, but this interval shrinks quickly starting in the mid 1980s as this probability falls. By the time of China's WTO accession in 2001, the bounds for this probability are only a few p.p. away from the point estimate. Whether one uses the point estimates reported in the main text or the bounds estimated in this appendix, the probability that the 1980 reform would be reversed was initially very high, began to fall rapidly in the mid 1980s, and was quite low in the years before and after WTO accession.

D.2 Unanticipated Changes in Trade-Policy Uncertainty

We have assumed that firms in our model know the entire path of probabilities of switching between trade policy regimes, i.e., that in 1971 firms know that the probability of losing NTR status will be high in the early 1980s but will be low later on. Here, 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.

[29] Recall that a higher probability of gaining MFN status pushes the gap elasticity during the 1970s upward, while a higher probability of losing MFN status pushes the gap elasticity downward after the 1980 reform.

Panel (b) of Figure F.8 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. This suggests that our approach provides tight bounds on these probabilities and on their economic effects.

D.3 Heterogeneous NTBs Instead of Aggregate TPU

We have assumed that changes in the elasticity of trade to the NNTR gap were driven by changes in the probability of switching from MFN tariffs to NNTR tariffs. This probability is the same across goods, so changes in the gap elasticity in our analysis are driven by heterogeneous responses across industries to trade policy uncertainty; high-gap industries are more sensitive to uncertainty than low-gap industries. Alternatively, one might ask: could the trajectory of the gap elasticity have been driven instead by heterogeneous changes across industries in other non-tariff barriers (NTBs)? Here, we explore how NTBs would have had to evolve across industries to rationalize our estimated gap elasticities in the absence of aggregate trade policy uncertainty.

Our approach is as follows. We assume that there is no uncertainty about trade policy; that is, tariffs follow the same deterministic trajectory as in the no-TPU counterfactual discussed in section 4.2. Instead, we add industry-specific, time-varying iceberg trade costs. Specifically, the firm's resource constraint is now

$$z_t \ell \geq d_{gt}(p, \tau_{gt}) \xi_t \psi_{gt}, \tag{D.1}$$

where ξ_t is the firm's idiosyncratic variable trade cost as before, and ψ_{gt} is a trade cost that is common to all firms in industry g (the NTB). To put some structure on how the industry-specific non-tariff barriers relate to the NNTR gap, we assume that

$$\psi_{gt} = \chi_t X_g, \tag{D.2}$$

where X_g denotes the gap in industry g as in the empirics, and χ_t is a parameter that governs how much higher these NTBs are in high-gap versus low-gap industries. Leaving all other aspects of the calibration unchanged, we choose χ_t so that the model matches the same smoothed path of the NNTR gap shown in panel (a) of figure 6.

Figure F.9 shows the results of this analysis. We find that χ_t is about -0.3 before 1980, jumps sharply by about 70 percentage points in 1980, and then gradually falls to zero following a similar path to the switching probability from our main quantitative analysis shown in panel (b) of figure 6. In other words, before the 1980 reform, high-gap industries actually had lower NTBs than low-gap industries, and then suddenly had much higher NTBs when this reform occurred. To make this concrete, we plot the mean NTB for industries in the top 25% of the NNTR gap distribution (labeled high-gap industries in the figure) alongside the mean for industries in the bottom 25% (labeled low-gap). Prior to the 1980 reform, the mean NTB for high-gap industries was about 15p.p. lower than the mean for low-gap industries. Immediately following the reform, the former was about 25p.p. higher than the latter. Thus, the 1980 reform increased—in the span of just one year—NTBs in high-gap industries by about 40p.p. relative to low-gap industries.

In our view, these results indicate that time-varying heterogeneity across industries in NTBs is unlikely to be the primary driver of the growth in the gap elasticity. To our knowledge, there is no evidence in the historical record or the data to suggest that the United States immediately and systematically replaced the high pre-1980 tariffs in high-gap industries with non-tariff barriers.

E Additional tables

Table E.1: ECM: additional controls and alternative samples

	Baseline	Shipping	All Goods	All Countries	Full	Balanced
$\mathbb{1}\{j = China\}\Delta\tau_{jgt}$	-2.29***	-2.22***	-2.63***	-2.31***	-2.66***	-2.35***
	(0.38)	(0.39)	(0.38)	(0.38)	(0.38)	(0.41)
$\mathbb{1}\{j = China\}\tau_{jg,t-1}$	-2.92***	-2.92***	-2.80***	-2.96***	-2.85***	-2.63***
	(0.30)	(0.30)	(0.26)	(0.29)	(0.26)	(0.27)
$\mathbb{1}\{j = China\}v_{jg,t-1}$	-0.37***	-0.37***	-0.37***	-0.37***	-0.37***	-0.31***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
$\mathbb{1}\{j \neq China\}\Delta\tau_{jgt}$	-1.91***	-1.81***	-1.98***	-2.08***	-2.18***	-1.97***
	(0.13)	(0.13)	(0.11)	(0.12)	(0.10)	(0.14)
$\mathbb{1}\{j \neq China\}\tau_{jg,t-1}$	-1.54***	-1.43***	-1.67***	-1.72***	-1.89***	-1.51***
	(0.11)	(0.11)	(0.09)	(0.09)	(0.08)	(0.12)
$\mathbb{1}\{j \neq China\}v_{jg,t-1}$	-0.47***	-0.48***	-0.47***	-0.46***	-0.46***	-0.44***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Shipping Costs$_{jgt}$		-2.78***				
		(0.05)				
Long-run China	−7.96***	−7.90***	−7.56***	−8.06***	−7.67***	−8.39***
	(0.87)	(0.88)	(0.75)	(0.87)	(0.75)	(0.88)
Long-run Others	−3.27***	−2.98***	−3.57***	−3.75***	−4.14***	−3.41***
	(0.87)	(0.88)	(0.75)	(0.87)	(0.75)	(0.88)
Long-/Short-Run China	3.48	3.56	2.87	3.49	2.88	3.57
Long-/Short-Run Others	1.71	1.65	1.80	1.80	1.90	1.73
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	939,788	939,788	1,100,988	1,053,446	1,232,395	572,759
R^2	0.36	0.37	0.36	0.35	0.35	0.34

Notes: All results are obtained estimating (2). 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 E.2: ECM: aggregation and time horizon

	1974-2008	1974-88		1989-2008	
	SITC	TS-USA	SITC	HS-8	SITC
$\mathbb{1}\{j = China\}\Delta\tau_{jgt}$	-2.29***	-1.85***	-2.32***	-2.37***	-2.20***
	(0.38)	(0.28)	(0.40)	(0.58)	(0.81)
$\mathbb{1}\{j = China\}\tau_{jg,t-1}$	-2.92***	-3.06***	-2.83***	-3.88***	-2.70***
	(0.30)	(0.27)	(0.39)	(0.43)	(0.88)
$\mathbb{1}\{j = China\}v_{jg,t-1}$	-0.37***	-0.75***	-0.61***	-0.51***	-0.45***
	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
$\mathbb{1}\{j \neq China\}\Delta\tau_{jgt}$	-1.91***	-2.36***	-2.12***	-4.02***	-1.66***
	(0.13)	(0.13)	(0.22)	(0.16)	(0.15)
$\mathbb{1}\{j \neq China\}\tau_{jg,t-1}$	-1.54***	-2.75***	-1.87***	-3.42***	-1.63***
	(0.11)	(0.15)	(0.21)	(0.15)	(0.15)
$\mathbb{1}\{j \neq China\}v_{jg,t-1}$	-0.47***	-0.77***	-0.66***	-0.60***	-0.55***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Long-Run China	-7.96***	-4.09***	-4.61***	-7.66***	-6.00***
	(0.57)	(0.36)	(0.41)	(0.78)	(1.51)
Long-Run Others	-3.27***	-3.55***	-2.84***	-5.71***	-2.98***
	(0.57)	(0.36)	(0.41)	(0.78)	(1.51)
Long-/Short-Run China	3.48	2.21	1.99	3.23	2.73
Long-/Short-Run Others	1.71	1.50	1.34	1.42	1.80
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	939,788	965,641	320,824	1,691,082	613,403
R^2	0.36	0.56	0.47	0.42	0.39

Note: All estimates are obtained by using (2). 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 1974–88, 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 including g are at the aggregation level of the sample used. Standard errors in parentheses are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table E.3: NNTR-gap elasticity: Pierce and Schott (2016) specification

	NNTR Gap		Statutory NNTR		Applied NNTR	
$\mathbb{1}\{^{t>2000}_{j=China}\}NNTRGap_g$	-0.92***	-2.50***				
	(0.22)	(0.28)				
$\mathbb{1}\{^{t>2000}_{j=China}\}NNTR_g$			-0.74***	-2.07***		
			(0.21)	(0.26)		
$\mathbb{1}\{^{t>2000}_{j=China}\}AppNNTR_g$					-0.85***	-2.71***
					(0.23)	(0.33)
τ_{jgt}	-3.28***	-3.20***	-3.28***	-3.21***	-3.04***	-3.24***
	(0.14)	(0.11)	(0.14)	(0.11)	(0.18)	(0.12)
Period	'92–'07	'74–'08	'92–'07	'74–'08	'92–'07	'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	781,960	1,453,687	387,442	839,364
R^2	0.86	0.81	0.86	0.81	0.88	0.82

Notes: All estimates are obtained from variations of (B.1). 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 (B.1). 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 E.4: NNTR-gap elasticity: robustness

	Baseline	Anticipation	Shipping	Tariffs	Life cycle
$\mathbb{1}\{^{t=t'}_{j=China}\}X_g$					
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		-0.16	-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 (4). 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—*Tariffs*—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 E.5: NNTR-gap elasticity: robustness, continued

	All	Full	Balanced	NNTR	Applied
$\mathbb{1}\{^{t=t'}_{j=China}\}X_g$					
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	0.01	-0.30	-0.20	0.15
2005	-0.64^{*}	0.06	-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 (4). 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 are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table E.6: NNTR-gap elasticity: TS-USA, 1974-88

	Baseline Sample	All Countries	Applied Tariffs Communists, 1974-79	Incl. Tariffs
$\mathbb{1}\{^{t=t'}_{j=China}\}X_g$				
1974	-4.27***	-4.23***	-3.93***	-2.99***
1975	-3.86***	-3.81***	-3.53***	-2.57***
1976	-3.73***	-3.69***	-3.44***	-2.43***
1977	-3.84***	-3.80***	-3.52***	-2.51***
1978	-3.33***	-3.27***	-2.88***	-2.03***
1979	-3.81***	-3.78***	-3.45***	-2.51***
1980	-2.66***	-2.61***	-2.21***	-2.45***
1981	-1.84***	-1.79***	-1.72***	-1.70***
1982	-1.99***	-1.93***	-1.82***	-1.94***
1983	-2.20***	-2.15***	-1.90***	-2.16***
1984	-2.11***	-2.05***	-1.79***	-2.08***
1985	-1.40***	-1.35***	-1.03***	-1.37***
1986	-0.98***	-0.93***	-0.66***	-0.95***
1987	-0.51**	-0.47**	-0.30	-0.50**
τ_{jgt}				-1.89***
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	486725	500641	701141	486725
Adjusted R^2	0.802	0.800	0.802	0.802

Note: All estimates are obtained by using (4), 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 **??** 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 (4) and as expected the coefficient diminishes in the early years. Standard errors are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

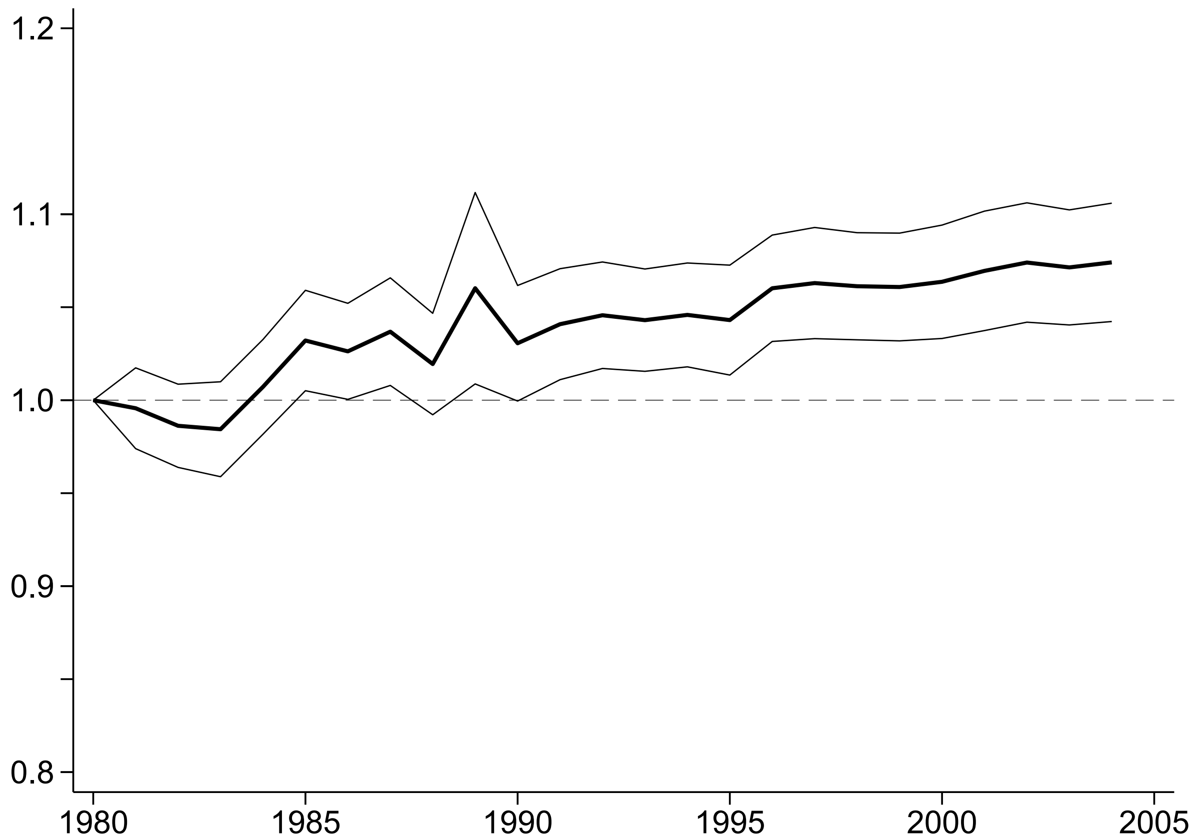
Table E.7: NNTR-gap elasticity: HS-8, 1990-2007

	Baseline Sample	No Tariffs	All Countries	Full Sample	Balanced g included if $v_{CHN,g,1990} > 0$
$\mathbb{1}\{^{t=t'}_{j=China}\}X_g$					
1989	-3.42***	-3.45***	-3.36***	-2.25***	-3.34***
1990	-2.67***	-2.70***	-2.60***	-1.82***	-2.58***
1991	-2.18***	-2.22***	-2.14***	-1.55***	-2.10***
1992	-1.59***	-1.64***	-1.55***	-1.09***	-1.54***
1993	-1.03***	-1.08***	-0.98***	-0.83***	-1.01***
1994	-1.20***	-1.26***	-1.18***	-1.07***	-1.14***
1995	-1.06***	-1.10***	-0.99***	-1.00***	-1.21***
1996	-0.84***	-0.87***	-0.80***	-0.84***	-0.80***
1997	-1.01***	-1.03***	-0.99***	-0.98***	-1.00***
1998	-0.99***	-1.00***	-0.99***	-1.11***	-0.88***
1999	-0.63***	-0.63***	-0.62***	-0.89***	-0.57**
2000	-0.46***	-0.47***	-0.48***	-0.88***	-0.31
2001	-0.50***	-0.52***	-0.51***	-0.99***	-0.29
2002	-0.29*	-0.30*	-0.30*	-0.60***	-0.15
2003	-0.32**	-0.32**	-0.35**	-0.56***	-0.33*
2004	-0.33**	-0.32**	-0.32**	-0.54***	-0.18
2005	-0.31**	-0.31**	-0.31**	-0.07	-0.09
2006	-0.28**	-0.28**	-0.31**	-0.14	-0.18
2007	-0.17	-0.16	-0.19	-0.02	0.05
τ_{jgt}	-3.87***		-3.37***	-4.07***	-4.62***
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	2,079,150	2,079,150	2,314,918	2930184	1,246,200
Adjusted R^2	0.766	0.766	0.774	0.774	0.778

Note: All estimates are obtained by using (4), 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 (4) (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 are clustered at the gj level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

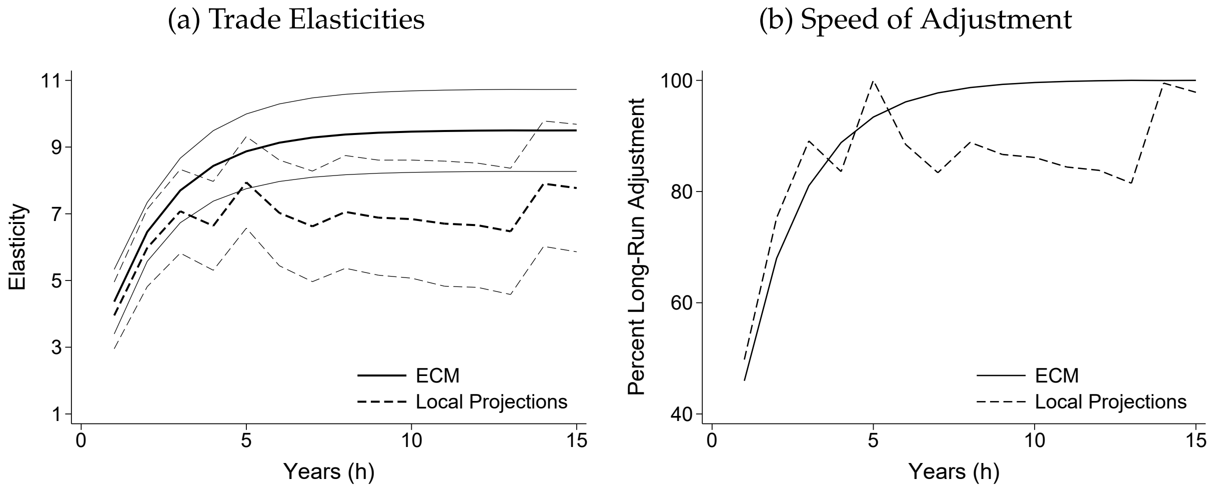
F Additional figures

Figure F.1: Autocorrelation of 1980 tariff changes



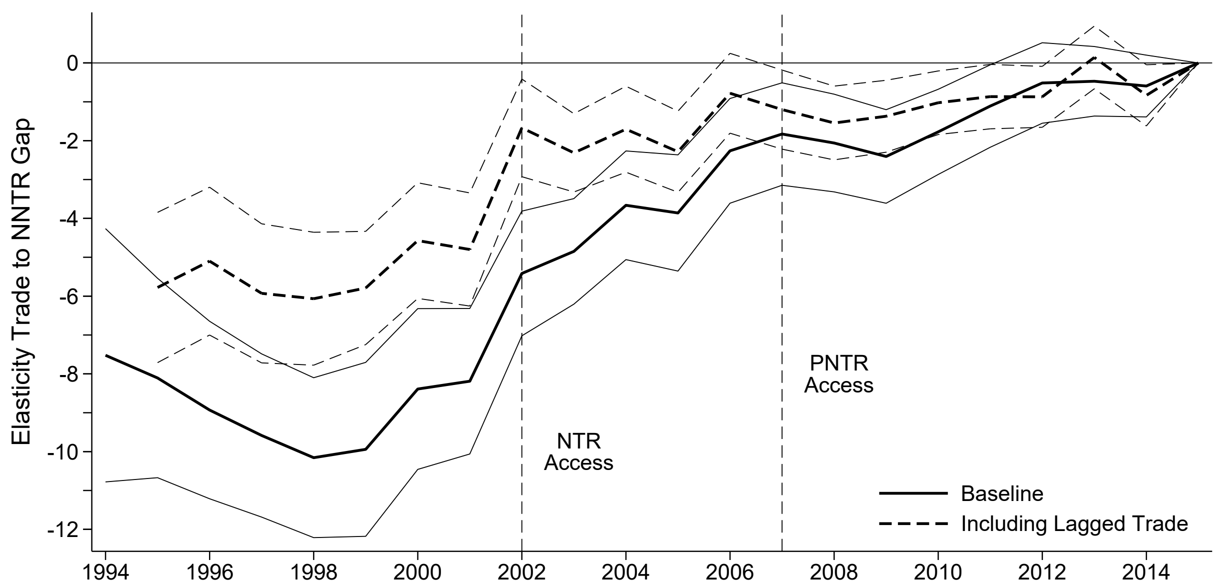
Notes: This figure plots the autocorrelation of the 1980 tariff changes, that are obtained from regressing the h-year ($h = [1, 25]$) tariff relative to 1979 on the one-year tariff change between 1980 and 1979. This regression includes the same fixed effects as in (3). The 95 percent confidence interval is estimated using standard errors clustered at the gj level.

Figure F.2: Vietnam's speed of adjustment to the 2002 NTR grant



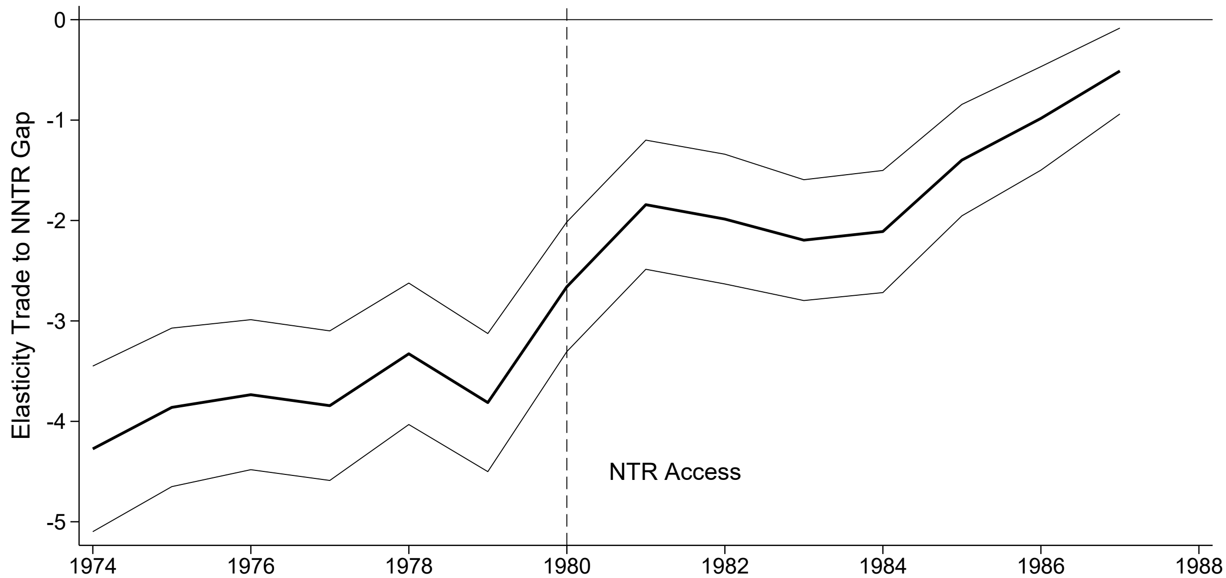
Notes: Panel (a) reports the time varying elasticities to tariff changes on U.S. imports from Vietnam. The solid line uses the ECM approach of (2) and the dashed line use the local projection approach of (3), focusing on changes after 2001, the year prior to Vietnam's NTR access. The standard errors that construct the 95 percent confidence intervals are clustered at the gj level. Panel (b) plots the ratio of the elasticity at each period relative to the maximum elasticity throughout all periods.

Figure F.3: Elasticity of U.S. imports from Vietnam to NNTR gap



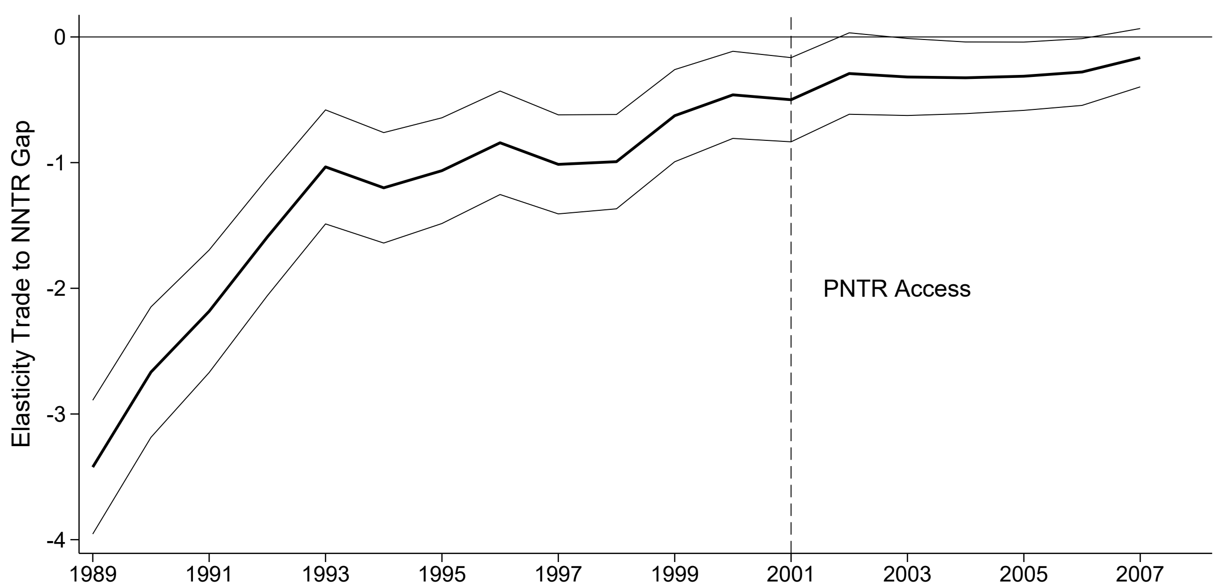
Notes: This Figure plots the estimates of $\hat{\beta}_t$ for $t = [1994, 2014]$ from (4) applied to Vietnam instead of China. The dashed additionally includes the lagged value of imports $v_{jg,t-1}$ on the right-hand side of (4). Note that the coefficient for 2014 is not identified because there are no lagged trade flows prior to the lifting of the embargo in 2014. The standard errors that construct the 95 percent confidence intervals are clustered at the gj level.

Figure F.4: NNTR-gap elasticity: TSUSA-8, 1974–87



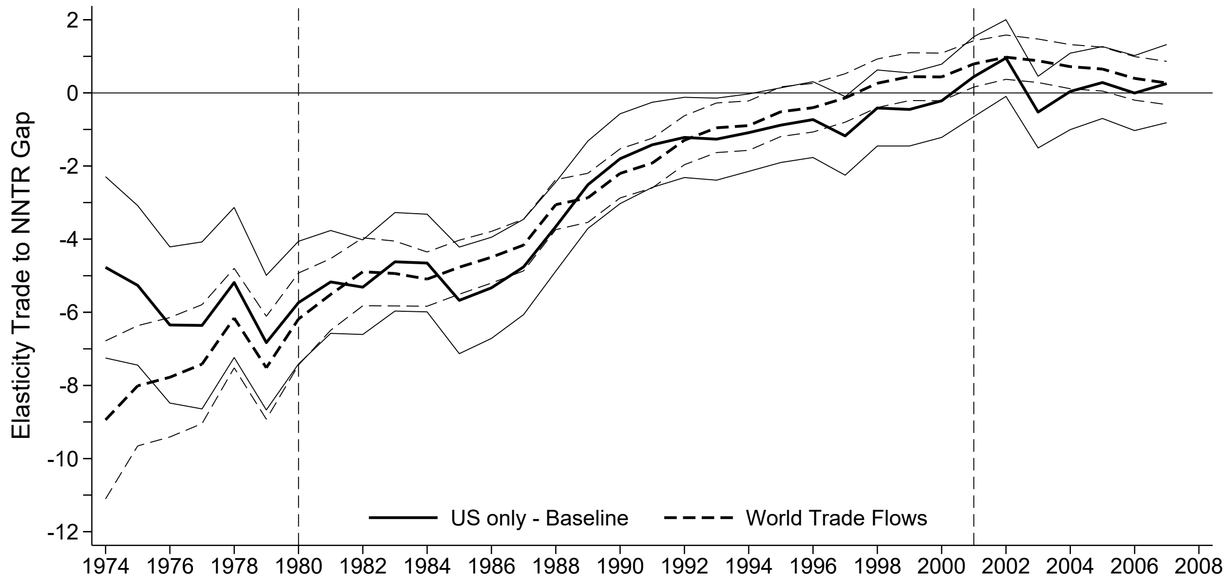
Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 1987]$ from (4) using 8-digit TSUSA product-level aggregation of all variables, instead of SITC. This level of aggregation ends in 1988. The standard errors that construct the 95 percent confidence interval are clustered at the gj level.

Figure F.5: NNTR-gap elasticity: HS-8, 1989–2007



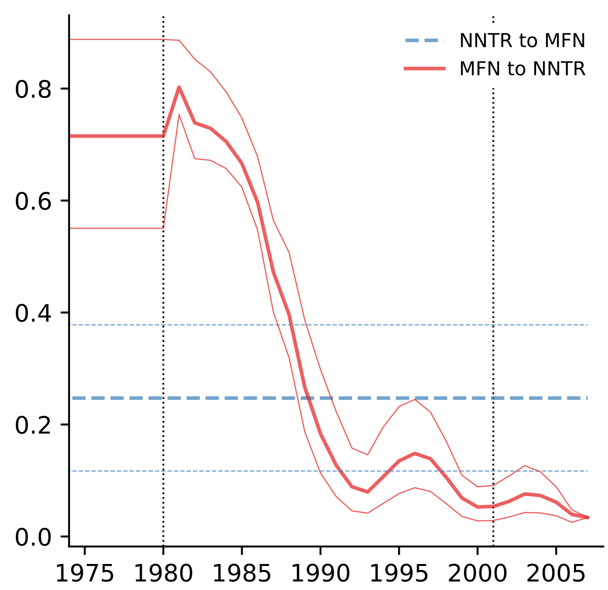
Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1989, 2007]$ from (4) using HS-8 product level aggregation of all variables, instead of SITC. This level of aggregation is only available since 1989. The standard errors that construct the 95 percent confidence interval are clustered at the gj level.

Figure F.6: NNTR-gap elasticity: China supply factors



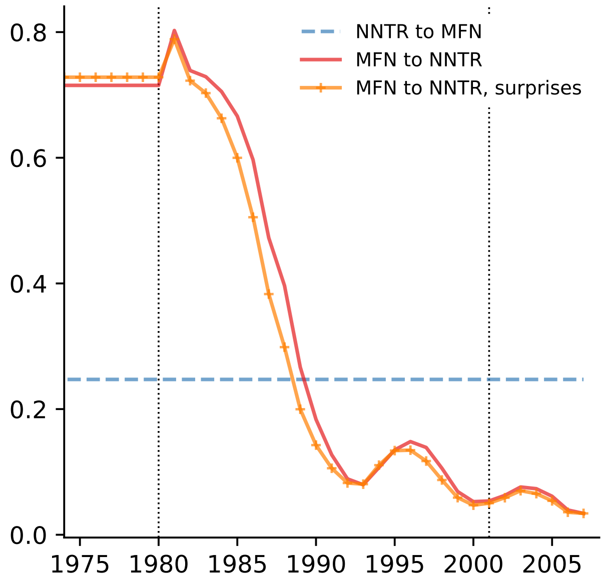
Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (B.2) using the merged World Trade dataset from Feenstra et al. (2005) (1974–2000) and the BACI Trade Dataset (2000–08). The standard errors that construct the 95 percent confidence interval are clustered at the gj level.

Figure F.7: Trade-policy probabilities: upper and lower bounds



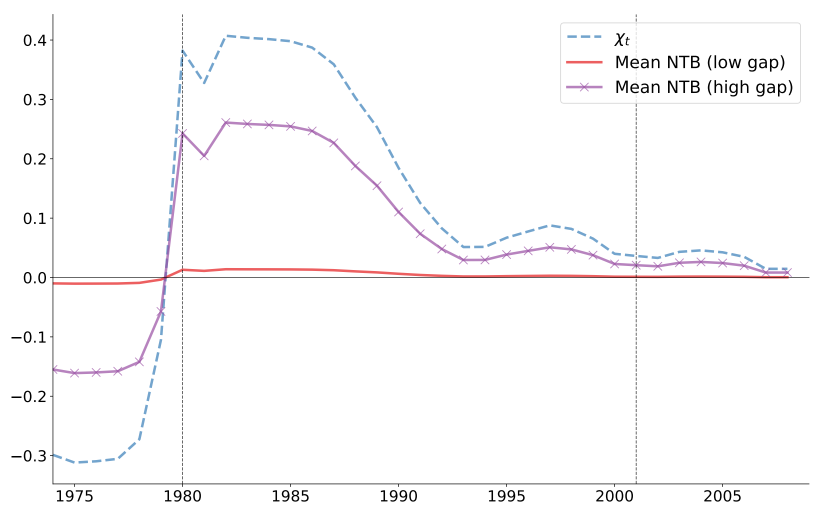
Notes: Shows the estimated probabilities of switching policy regimes. Thick lines are the baseline results. Thin lines are estimated by matching the upper and lower confidence intervals shown in Figure 4.

Figure F.8: Trade-policy probabilities: unanticipated changes



Notes: Shows the estimated probabilities of switching policy regimes. In surprises model, every year firms believe that current transition probabilities will apply forever; changes in transition probabilities are treated as unanticipated shocks.

Figure F.9: Time-varying non-tariff barriers instead of uncertainty



Notes: Shows the estimated industry-specific iceberg trade costs described in Appendix D.3. χ_t is the correlation between the NNTR gap and iceberb trade costs. Low-gap goods have NNTR gaps below the 25th percentile. High-gap goods have NNTR gaps above the 75th percentile.