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 embargo during 1950–1970 to 15 percent of overall U.S. imports in 2008, exploiting the rich heterogeneity across products in trade policy and trade growth. 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 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. However, separately identifying these two effects using data alone is difficult. We disentangle these effects using a structural model to estimate a path of trade-policy expectations. We find the 1980 reform was largely a surprise and initially had a high chance 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 while largely ignoring the effects of changes in trade policy that occurred in the past or may occur in the future. One recent literature shows that trade responds gradually to past policy changes, whereas another shows that uncertainty about future policy can affect trade in the present. We argue the effects of past and future policy are tightly intertwined, develop a method to disentangle these effects, and use it to study the growth of Chinese exports to the United States.

Our objective is to estimate how expectations of U.S. trade policy toward China evolved over time and measure how this evolution, together with observed changes in tariffs, shaped the growth of Chinese exports. Different from previous studies that focus on the 1990s and 2000s, we take a longer-term view that begins when the United States first opened to trade with China in 1971. We find that during the 1970s, when the United States levied high Non-Normal Trade Relations (NNTR) tariffs on Chinese exports, the probability that China would gain Normal Trade Relations (NTR), and face lower NTR tariffs, was about 30 percent.[1] This substantially increased exports of goods with large gaps between the two tariff schedules, which we refer to as high-gap goods. After formal diplomatic relations were established in 1979 and China gained NTR status in 1980, there was a high probability that this status would be revoked, which depressed exports of these same high-gap goods. This reform became more credible during the late 1980s and early 1990s, and this change in expectations was the main driver of sustained growth in Chinese exports to the United States, even after China joined the World Trade Organization (WTO) in 2001.

Our most important findings are that trade-policy uncertainty played a larger role in Chinese export growth during the 1980s than during the 1990s and 2000s, and that China's 2001 WTO accession had little impact on policy expectations or trade flows. Thus, we offer a different narrative than studies such as Pierce and Schott (2016) and Handley and Limão (2017), who interpret the fact that exports of high-gap goods grew faster than exports of low-

[1] Before 1998, the United States used the term Most Favored Nation (MFN) rather than Normal Trade Relations. We use the terms NTR and MFN interchangeably.

gap goods following WTO accession as evidence that this event significantly reduced policy uncertainty. We offer two new perspectives on this evidence. First, post-WTO high-gap export growth wasn't actually that rapid: high-gap exports grew six times faster relative to low-gap exports during 1985–1993 than during 2000–2006. Second, we find that much of the observed growth during the later period was a gradual adjustment to the reductions in policy uncertainty that occurred during the earlier period, and to the tariff liberalizations in 1971 and 1980. This new narrative highlights the importance of accounting for the interaction between expectations about future reforms and gradual adjustments to past reforms and changes in expectations.

Two aspects of U.S. trade policy towards China make it ideally suited to study the dynamics of trade policy uncertainty. First, as recognized by Handley and Limão (2017) and Pierce and Schott (2016), the United States maintains a dual tariff scheme wherein all of a country's exports are subject to the same tariff schedule (NTR or NNTR) at a given point in time, but tariffs under each schedule differ across goods. Thus, the possibility of switching between schedules affects some goods—those with large gaps between NTR and NNTR tariffs—more than others. Second, the United States imposed a trade embargo on China from 1950 to 1971. Every good started with zero trade, eliminating concerns about pretrends and allowing us to study disaggregated trade flows over the entire trade relationship.

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. 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., 2023), we find the long-run response of Chinese exports to U.S. tariffs is almost four times as large as the short-run response and that completing 90 percent of the adjustment takes as long as 20 years. The second measurement is an estimate of how trade responds to trade policy uncertainty (TPU). Building on Pierce and Schott (2016), we measure the response of trade in goods with large differences between NNTR and NTR 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. The larger the chance of losing this status, the lower exports of high-gap goods will be relative to exports of low-gap goods. We measure the elasticity of Chinese exports to the NTR gap for each year from 1974 to 2007. These results indicate that effects of policy uncertainty were the largest in the 1970s and early 1980s, and these effects had largely dissipated before China joined the WTO in 2001.[2]

Alone, neither of these measurements identify the effects of past policy or uncertainty about future policy, because the NTR gap is almost perfectly correlated with the change in tariffs in 1980. Our elasticities of trade to past policy are biased by uncertainty about future changes. If the risk of losing NTR status depressed trade prior to China's WTO accession, the measured response of trade to the initial 1980 liberalization would be smaller than in the absence of this risk. The NTR-gap elasticities are biased by the slow adjustment to this liberalization. Goods with the largest NTR gaps are most sensitive to the risk of losing NTR status, but these same goods also experienced the largest tariff reductions when this status was first granted, and thus took the longest to converge to their new levels of trade. We overcome these identification challenges by using indirect inference to estimate the time-varying transition probabilities between these two policy regimes in 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. (2021). It 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). Firms in each industry differ in terms of productivity, as well as variable export costs, which they reduce gradually through a risky investment.[3] The export entry decision and gradual reduction in export costs 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 NTR, and the probability of switching between regimes varies over time. This uncertainty depresses export participation and reduces trade volumes when the economy is in the NTR regime.

The model is calibrated using indirect inference to match our estimates of the adjustment

[2]That the effect of the NTR gap on trade was largest during the 1970s, before China gained NTR status, suggests the NTR gap captures something besides exposure to tariff risk. This possibility is one focus of this paper.

[3]Drozd and Nosal (2012), Fitzgerald et al. (2023), Piveteau (2021), and Steinberg (2021) develop similar models of slow firm-level adjustment to market entry through the accumulation of customers.

process and the year-by-year elasticities of trade to the NTR gap. Firms in the model understand the tariff regime can change, 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 NTR regime in 1980. The probabilities of switching between trade-policy regimes are chosen so that the transition to the steady state replicates our estimated path of the elasticities of trade to the NTR gap. Our identification works as follows. A higher likelihood of reverting from NTR tariffs to NNTR tariffs raises the expected 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.

The main output of our calibration is a time-varying path of transition probabilities between NTR and NNTR status. We find the annual probability of China gaining NTR status during the 1970s was about 30 percent. Once China gained NTR status in 1980, the probability of losing this status was initially high, peaking at 62 percent in 1981. This observation reflects our empirical finding that trade in high-gap goods stagnated relative to trade in other goods for several years after the 1980 reform. Starting in 1986, when China applied to join the international trade arrangement that would become the WTO, 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 Tiananmen 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 NTR tariff rates; this probability fell by less than 2 percentage points between 1999 and 2007.

We use our model to isolate the roles of trade-policy uncertainty and gradual adjustment to tariff changes by conducting a counterfactual in which there is no uncertainty about trade policy. In this counterfactual, aggregate trade grows faster, particularly in high-gap industries, so the elasticity of trade to the NTR gap shrinks quicker than in the benchmark model (and in the data). However, the gradual adjustment to the liberalizations in the counterfactual is still ongoing even after China's 2001 WTO accession. We find that gradual adjustment accounts for almost one-third of the overall change in the NTR-gap elasticity documented by Pierce and Schott (2016), which the literature has attributed entirely to a reduction in TPU caused by PNTR access.

Our model highlights a subtle, yet important, aspect of trade-adjustment dynamics: trade adjusts slowly to changes in expectations about future policy as well as past policy changes. To illustrate this point, we construct a counterfactual model in which the probability of losing NTR status remained constant until China joined the WTO. The single estimated probability of losing NTR status—and the extent to which WTO accession reduced this probability—is substantially higher in this experiment than in our benchmark model. This demonstrates that time-varying policy uncertainty is a key factor in explaining the path of China's integration into the U.S. market. Much of the growth in exports of goods with high NTR gaps after the turn of the century was a delayed adjustment to the increase in the credibility of U.S. policy toward China during the 1980s and 1990s, rather than a reduction in tariff risk associated with WTO accession.

In the last part of the paper, we revisit the effects of trade policy uncertainty on U.S. labor markets. Studies like Pierce and Schott (2016) document large declines in employment in U.S. industries with high NTR gaps after China joined the WTO. On the surface, these findings seem at odds with our new narrative. Using a simple model, we illustrate how differential growth of imports across products from China caused by changes in trade policy (or expectations about future policy) affects domestic employment across industries. Two key channels modulate this effect: the share of domestic absorption that is imported from China (import exposure); and the share of U.S. domestic production that is exported to world markets (export exposure). The same change in imports has a larger effect on employment in industries that are initially more exposed to Chinese imports and less exposed to exporting. We develop an estimating equation in which the effect of the NTR gap is interacted with these exposure terms. Like Pierce and Schott (2016), we find that U.S. employment has indeed declined as a result of changes in expectations about U.S. trade policy towards China. In contrast to their findings, we find most of this effect occurred long before China joined the WTO, consistent with our analysis of Chinese exports. When we restrict the effect of the NTR gap on employment to be the same across industries, we recover the same patterns as Pierce and Schott (2016). Accounting for heterogeneity in import and export exposure is crucial to getting the timing of the employment effects right.

This paper contributes to two strands of literature. The first studies the dynamics of trade after changes in trade policy. Baier and Bergstrand (2007) and Baier et al. (2014) show 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, Anderson and Yotov (2020), Khan and Khederlarian (2021), and Boehm et al. (2023) estimate long-run tariff elasticities of trade that are three to four times the short-run elasticities. Many of these studies grapple with issues related to endogeneity and anticipation of phased-in or temporary tariff changes. Our contribution is to study the dynamic response to a single exogenous, immediate tariff reduction over a multi-decade span using disaggregated data. We document similar results using two 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., 2023).

The second strand studies how expected future trade reforms 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.[4] Pierce and Schott (2016), Feng et al. (2017), Handley and Limão (2017), and Bianconi et al. (2021) measure the growth in trade (or other outcomes, such as employment or stock returns) that resulted from the elimination of uncertainty that occurred when China was granted PNTR status in 2001. 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. More generally, our work is related to the classic "Peso problem" of estimating the likelihood of discrete events that are possible but unobserved for long periods of time.

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

[4]Additionally, several papers study the impact of uncertainty about Brexit, such as Steinberg (2019), Crowley et al. (2019), and Graziano et al. (2021).

to estimate the probability of NNTR reversal. We estimate a time-varying probability over a longer interval using a richer model of exporter dynamics in which changes in tariffs have persistent effects. Our analysis highlights the importance of earlier changes in trade-policy expectations during the late 1980s and early 1990s in explaining the growth of Chinese exports to the United States around the turn of the century. We find that ignoring these changes leads to an overstatement of 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 and presents the results of our empirical analyses. Section 3 lays out our model and Section 4 discusses our calibration strategy. Section 5 presents the results of our quantitative analysis. Section 6 relates our findings about export growth to previous findings about the impact of this growth on U.S. labor markets. Section 7 concludes.

2 An empirical history of U.S.-China trade

We take a two-pronged approach to analyzing the joint dynamics of U.S. trade policy toward 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 the 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. Neither of these approaches can disentangle the effects of slow adjustment to past policy changes from uncertainty about future policy, but they provide strong evidence that these margins are important and 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.[5] We refer to this level of aggregation as a *good* and denote it by g. Our sample contains 1,742 goods, and includes 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 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 U.S. imports from China and every other country 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.[6] A key feature of our sample is that all countries, including China, faced approximately the same tariffs after 1980.[7] Thus, most tariff changes in our sample are explained by good-year fixed effects, and the main source of variation in tariffs applied to China is the 1980 liberalization. Moreover, the inclusion of imports by non-China NTR countries allows us to control for good-specific U.S. demand shocks. We also exclude goods that include products covered by the Multi Fiber Arrangement (MFA). As documented by Bambrilla et al. (2010), China's accession to the WTO triggered the removal of import quotas on these goods, and our tariff

[5]More disaggregated product-level data, such as 8-digit TSUSA 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 TSUSA and HS data during the periods in which these data are available.

[6] 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.

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

measures do not capture this material change in applied trade policy. In the appendix, we show our results are robust to alternative sample designs.

2.2 Policy dynamics

Between the imposition of the trade embargo in December of 1950 and 2008, there were three key reforms to U.S. import tariffs on Chinese goods: 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 import tariffs 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 for the fifteen sectors at the 2-digit level of the Chinese Industrial Classification System.[8] 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 toward China relative to countries that had NTR status throughout our sample period. Figure 2(a) plots the inverse tariff on

[8] We use this classification system to be consistent with the Chinese firm-level data that we use to calibrate the model in Section 4. We concord it to the SITC classification system using the ISIC Revision 4 concordance. The results are similar when we report these statistics at the 1-digit SITC level.

Chinese goods normalized by the tariff rate applied to other NTR countries:

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

In 1971, when the embargo ended, the inverse 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 tariff cuts were both immediate. Figure 2(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. Some small variation occurs from 1981 onward due to specific tariffs and temporary commercial policy (as well as aggregation issues and measurement error), but this variation is minor relative to the effects of NTR access.

In addition to the changes in tariffs, U.S. trade policy on imports from China was uncertain, and the extent of that uncertainty varied over time with domestic and international politics.[9] Following the lifting of the embargo in 1971 and gaining access to the NNTR rates, it was unclear what path would lead to granting China MFN rates. A series of U.S. trade acts in 1951, 1962, and 1974 required imports from non-market economies to face the higher NNTR rates. Exceptions to this requirement were rare, typically contingent on substantial reforms, and could be removed quickly (or withdrawn before being implemented). There were, however, periodic discussions of the wide-scale removal of this requirement on non-market economies.[10] A further source of uncertainty arose from the lack of diplomatic relations between the United States and the People's Republic of China in the 1970s. Owing to the One China principle, recognition of the People's Republic of China required with-

[9]Uncertainty about U.S. trade policy towards China precedes the embargo, owing to the Chinese Civil War and U.S. support for the Nationalists. Uncertainty about trade policy in this earlier period is beyond the scope of the paper.

[10]In 1966, the Johnson administration proposed the East-West Trade Relations Act that would provide NTR rates to the Soviet Union and other Communist countries of Eastern Europe. Following the 1972 U.S.-U.S.S.R. Trade Agreement, congressional debate led to the Jackson-Vanik amendment that linked freedom of emigration to MFN status in 1974 Trade Act. The Soviet Union was not granted MFN status until 1992. In 1977 the U.S. International Trade Commission made reports on the economic effects of providing access to U.S. markets to China and the Soviet Union at MFN rates.

drawing support for the Republic of China (Taiwan) and the mutual defense treaty. When President Carter did this in December of 1978, there was much debate about its legality and Congress swiftly proposed, and overwhelmingly passed, the Taiwan Resolution Act in 1979. Carter's likely challenger, and ultimately the next President, Ronald Reagan, seized on this issue, emphasizing that dropping support for Taiwan set a bad precedent for other countries.[11]

When China was granted temporary NTR status in 1980, it was uncertain how long this status would last. The program to allow non-market economies access to U.S. markets at MFN rates was contingent on satisfying conditions of the Jackson-Vanik Amendment to the Trade Act of 1974. China was the third country to gain access to MFN tariffs under this program, following Romania in 1975 and Hungary in 1978. After China, no country gained access for another 10 years. In the meantime, Romania lost NTR status from 1988–1991, Poland from 1982–1987, and Serbia and Montenegro from 1991–1992. Every year, the U.S. president had to renew China's NTR status by July, subject to Congressional approval. The first renewal was granted by President Reagan.[12] Each year from 1990 onward, the U.S. Congress voted to approve or disapprove this renewal. 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 NTR (PNTR) status contingent on joining the WTO. The PNTR status grant came into effect in December 2001 when China officially entered the WTO. The Trump tariffs of 2018, however, make it clear that gaining PNTR status did not entirely eliminate the risk of further tariff changes.

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. Trade is known to adjust gradually to trade liberalizations. For example, Baier and Bergstrand (2007) find that trade doubles in 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

[11]See, "Decency for Taiwan" (Reagan, 1979).

[12]The change in administration in 1981 likely increased uncertainty about China's NTR status because President Ronald Reagan was more openly anti-communist than President Carter and was very supportive of Taiwan. This is precisely what we find in our quantitative analysis.

al. (2023), 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 the growth in trade that followed the 1980 reform was gradual and took many years to complete.

Our first approach is an 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 determined 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}) capture 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.[13] Because our sample excludes countries, other than China, that did not receive NTR status during the sample, the good-year fixed effects absorb the effects of multilateral changes in MFN tariff rates on trade; thus, σ^{SR} and σ^{LR} measure how U.S. imports responded to changes in bilateral tariffs above and beyond these multilateral changes. For the China-specific elasticities σ^{SR}_{China} and σ^{LR}_{China}, given that the vast majority of the identifying tariff variation occurs during 1979-1981, these elasticities largely capture the way that Chinese exports responded over time to

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

the 1980 NTR grant.

The solid line in Figure 3(a) shows the path of adjustment to a one-time tariff change implied by our ECM estimates; the estimated parameter values are reported in Table F.2 of the appendix. The short-run trade elasticity is -2.29 and the long-run elasticity is -7.96. The former is consistent with many other estimates in the literature, and the latter, although large, is similar to the documented effects of other major liberalizations.[14] 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 gradual. As can be seen in Figure 3(b), our estimates imply U.S. imports from China take seven years to complete 90 percent of the total long-run adjustment to a tariff change. We show in Appendix B 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: $v_{jg,1979+h} - v_{jg,1979}$, for $h = 1, 2, \ldots, 25$.[15] We follow Boehm et al. (2023) and instrument the h-year change in tariffs relative to 1979 with the tariff change between 1980 and 1979 to account for the autocorrelation of this tariff change.[16] The fixed-effect structure is the same as in (2), except δ_{jg} is eliminated by taking differences of the dependent variable.

The dashed line in Figure 3(a) shows the path of adjustment to a tariff change implied by our local-projections estimation. Whereas the short- and long-run elasticities obtained by

[14] Khan and Khederlarian (2021) and Alessandria et al. (2023a) estimate similar elasticities for Canadian and Mexican exports to the United States following the creation of NAFTA, and for Vietnamese exports to the United States after Vietnam was granted NTR status in 2002. Yilmazkuday (2019) reports similar numbers from a VAR approach.

[15] In Alessandria et al. (2023a), we show 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 only changes relative to 1979.

[16] Figure G.1, in the appendix, shows the tariff changes from the NTR access were permanent and slightly increased over time. This observation contrasts with the predominantly mean-reverting tariff changes over the full sample.

the local projections are very close to those from the ECM, the non-parametric estimation of the year-by-year adjustment displays a much slower transition (Figure 3(b)). Our local-projections estimates imply completing 50 percent of the total adjustment takes 10 years and completing 90 percent takes 20 years. The final adjustment (although statistically insignificant) occurs after 22 years, precisely when China joined the WTO and gained PNTR status. Appendix B shows these results, too, are robust to a wide range of alternative specifications.

Overall, these results indicate 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 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, changes in expectations about future tariffs likely occurred. If the 1980 reform was initially viewed as unlikely to be permanent, as we find in our quantitative analysis, then the initial trade response to this reform would have been smaller, and the adjustment process longer, than it would have been in the absence of uncertainty.

2.4 Effects of the risk of losing NTR status

Our second empirical approach draws from the TPU literature, particularly the seminal studies of Pierce and Schott (2016) and Handley and Limão (2017). These studies document that the growth in U.S. imports of products from China around China's WTO accession was strongly correlated with the gap between the NNTR and NTR rate, 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 observation is taken as evidence that the risk of future tariff increases depressed trade in those products and that eliminating this risk stimulated trade growth in these products. 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 granted NTR status in 1980.

The unifying theme in the TPU literature is a difference-in-differences empirical strategy that compares goods that were more exposed to uncertainty about future policy to goods that were less exposed. In the U.S.-China context, exposure to policy uncertainty is typically measured by the NTR 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 NTR gap in 1999 as a time-invariant measure of the NTR gap: $GAP_g = \log(1 + \tau_g^{NNTR} - \tau_{g,1999}^{MFN})$. 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 NTR 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 NTR gap*, or the *NTR-gap elasticity*.

Figure 4 shows our results. Between 1974 and 1979, the NTR-gap elasticity was relatively stable around -10, indicating 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 NTR 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 NTR-gap elasticity during this period was large 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 the NTR gap in 1999 is highly correlated with the changes in applied tariffs on Chinese goods that occurred between 1979 and 1981.[17] Note this observation also explains why the NTR-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 NTR-gap elasticity rose sharply and then leveled off for several years. It did not begin to grow steadily until 1986. Our interpretation of

[17] A linear regression yields a coefficient of 0.82 and an R-squared of 0.88. The correlation would be perfect if not for the modest reductions in MFN tariffs that have occurred since 1981.

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 sustained growth in high-gap imports, despite 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 NTR gaps fell during this period.

The NTR-gap elasticity again leveled off from 1992 until 1998, before rising 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. The effect of this event on the NTR-gap elasticity, however, was relatively small in comparison to the changes that occurred during the 1980s. This finding suggests 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. Our estimates show the change in the NTR-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 NTR-gap elasticities using more aggregated global data on bilateral imports and exports (not only U.S. imports) to consider the role of good-specific Chinese supply factors. This approach allows us to control for any spurious correlation between the NTR gap and differences across goods in, for example, export licenses, state-owned enterprises, import quotas, and spillovers from infrastructure development. Appendix C 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 of the effect of this risk are likely to be confounded by the slow adjustment process. The NTR 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 NTR gap captures the lagged effect of the NTR status grant is likely to diminish in time.

3 Model of trade-policy dynamics

We use a structural model to isolate the roles of gradual adjustment and policy uncertainty on the growth of U.S. imports from China. Our model builds on Handley and Limão (2017) and Alessandria et al. (2021). There are G goods that correspond to the 5-digit SITC goods in our empirical analysis. Within each good g, a continuum of heterogeneous Chinese 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 is fixed: when a firm that produces good g 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. There are two trade-policy regimes, NNTR and MFN. The probability of switching between regimes varies over time, generating time-varying tariff risk.

Trade policy. Tariffs on each good, $\tau_{gt}(s)$, depend on the current trade-policy regime, s, which is an aggregate state variable that can take two values: MFN ($s = 1$), or NNTR ($s = 2$).[18] The tariff regime follows a time-varying Markov process with transition probabilities $\omega_t(s, s')$. For example, if the current regime is NNTR, the probability of being granted MFN status in the next period is $\omega_t(2, 1)$. We assume that firms know the entire path of regime-switching probabilities, $\{\omega_t(s, s')\}_{t=0}^{\infty}$. We consider alternative information structures in Section 5 and Appendix D.

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

$$y = z\ell. \tag{5}$$

Productivity, z, is independent across firms and follows a good-specific, stationary Markov process with transition probabilities $h_g(z, z')$. U.S. demand for a firm's good, d_{gt}, is a downward-sloping function of the tariff on that good and the price the firm charges, p,

$$d_{gt}(p, s) = (p\tau_{gt}(s))^{-\theta_g} D_{gt}. \tag{6}$$

[18] We abstract from good-specific risks such as anti-dumping duties or other good-specific commercial policy.

D_{gt} is an aggregate demand shifter and θ_g is the good-specific price elasticity of demand. Note that the aggregate trade elasticity is determined by the export participation response to a tariff change as well as the demand elasticity.

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 ($\infty > \xi_{gH} > \xi_{gL}$) 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. When a firm that did not export in the previous period chooses to export, it begins with ξ_{gH}. Following Alessandria et al. (2021), we assume the probability a firm retains its current ξ is symmetric: $\mathrm{P}(\xi_{gL}|\xi_{gL}) = \mathrm{P}(\xi_{gH}|\xi_{gH}) = \rho_\xi$. This specification implies exporters start exporting a small share of sales and, with some luck and repeated investments, grow to a larger export intensity.

The second type of trade cost is a fixed cost, f, that the firm must pay to export in the next period. The fixed costs are identical across firms within a good, but are a function of the firm's export history. If the firm is currently a nonexporter, it pays f_{g0} to enter the export market next period. If the firm is currently exporting, it pays f_{g1} to continue exporting in the next period. We summarize the fixed-cost structure in a function, $f_g(\xi)$, where $f_g(\infty) = f_{g0}$ and $f_g(\xi_{gL}) = f_g(\xi_{gH}) = f_{g1}$. This model generalizes the sunk-cost model of Das et al. (2007) in a way that captures the exporter life cycle (Ruhl and Willis, 2017).

We allow the trade costs (both fixed and variable) to differ across goods, which will help us capture heterogeneity in export participation dynamics across sectors in our calibration. The probability of switching variable trade costs (ρ_ξ) is assumed to be constant across goods, as this parameter primarily governs the aggregate long-run trade elasticity.

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,

$$\pi_{gt}(z, \xi, s) = \max_{p,\, \ell} p\, d_{gt}(p, \tau_{gt}(s)) - w_t \ell \tag{7}$$

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

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

$$V^1_{gt}\left(z, \xi, s\right) = -f_g(\xi) + \frac{\delta(z)}{1+r} \sum_{s'} \omega_t(s, s') \mathop{\mathbb{E}_t}_{z', \xi'} V_{g,t+1}\left(z', \xi', s'\right), \tag{9}$$

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

$$V^0_{gt}\left(z, \xi, s\right) = \frac{\delta(z)}{1+r} \sum_{s'} \omega_t(s, s') \mathop{\mathbb{E}_t}_{z'} V_{t+1}\left(z', \infty, s'\right), \tag{10}$$

and the value of the firm is

$$V_{gt}\left(z, \xi, s\right) = \pi_{gt}(z, \xi, s) + \max\left\{ V^1_{gt}\left(z, \xi, s\right), V^0_{gt}\left(z, \xi, s\right) \right\}. \tag{11}$$

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

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

which can be rewritten as

$$f_g(\xi) = \frac{\delta(\bar{z}_{gt}(\xi, s))}{1+r} \sum_{s'} \omega_t(s, s') \left\{ \mathop{\mathbb{E}_t}_{z', \xi'} \left[V_{t+1}(z', \xi', s') \right] - \mathop{\mathbb{E}_t}_{z'} \left[V_{t+1}(z', \infty, s') \right] \right\}. \tag{13}$$

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, the gain in firm value depends on the entire expected path of future tariffs, not the current applied tariff rate.

Aggregation. The decision rules, $\bar{z}_{gt}(\xi, s)$, determine how the distribution of productivity and variable trade costs across firms, $\varphi_{gt}(z,\xi)$, evolves over time for a sequence of realizations of the aggregate state, $\{s_t\}_{t=0}^{\infty}$. The law of motion for this distribution is given by

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

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

$$\begin{aligned}
\varphi_{g,t+1}(\mathcal{Z},\xi_{gL}) = & (1-\rho_\xi)\int_{\bar{z}_{gt}(\xi_{gH},s_t)}^{\infty} h_g(\mathcal{Z},z)\varphi_{gt}(z,\xi_{gH})dz \\
& + \rho_\xi \int_{\bar{z}_{gt}(\xi_{gL},s_t)}^{\infty} h_g(\mathcal{Z},z)\varphi_{gt}(z,\xi_{gL})dz,
\end{aligned} \tag{16}$$

where $\mathcal{Z}$ is a typical subset of $\mathbb{R}_{++}$, $h_g(\mathcal{Z},z)$ is the probability of surviving and drawing a new productivity in $\mathcal{Z}$ conditional on today's productivity z, and $\bar{h}_g(\mathcal{Z})$ is the probability of dying and being replaced by a new firm with productivity in $\mathcal{Z}$. Although the decision rules, $\bar{z}_{gt}(\xi,s)$, respond immediately to trade-policy changes, the export participation rate adjusts gradually as new firms draw productivity shocks that are good enough 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,

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

respond slowly to policy changes. The slow adjustment 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.

4 Calibration

Our calibration strategy has four main stages. First, we group our 5-digit SITC goods into 15 broad sectors and use Chinese firm-level panel data to compute statistics about exporter dynamics for each of these sectors during 2004–2007. Second, we assign several parameters to standard values from the literature and make some common functional form assumptions. Third, we calibrate each sector's productivity dispersion and non-tariff exporting costs so that the model's terminal steady state replicates these statistics. Fourth, we calibrate the probabilities of switching iceberg costs and trade-policy regimes so that the model's transition matches our estimates of the long-run trade elasticity in Section 2.3 and the time-varying NTR-gap elasticity in Section 2.4.

4.1 Chinese firm-level data and sectoral heterogeneity

We follow the trade-dynamics literature and calibrate our model to match micro-level facts about exporter life cycles and macro-level facts about trade dynamics. The micro-level facts come from an annual survey of manufacturing enterprises collected by the Chinese National Bureau of Statistics.[19] We use this data to compute four statistics that summarize the distribution and dynamics of Chinese firms that export to the United States: the dispersion in export sales (as measured by the coefficient of variation of log exports); the fraction of firms that export (the export participation rate); the fraction of exporters who stop exporting each period (the exit rate); and the average exports of an incumbent exporter compared to the average exports of a new exporter (the incumbent premium).

A potential concern with calibrating our model to match these statistics is aggregation bias. Specifically, exporter dynamics could vary across sectors, reflecting sectoral heterogeneity in non-tariff trade costs and other technological primitives. If this heterogeneity was correlated with the NTR gap, the evolution of the elasticity of trade to this gap, shown in Figure 4, could be driven by changes in the sectoral composition of Chinese exports to the United States rather than changes in expectations about U.S. trade policy towards China.[20] To ensure that our results are not driven by aggregation bias, we compute the statistics described above for each of the fifteen sectors in the 2-digit China Industry Classification System. We map the 5-digit SITC goods in our data to this system using the ISIC Revision 4 concordance. Table 3 lists the fifteen sectors and reports the four moments for each. There

[19] This data has been widely used to study Chinese manufacturing growth between the late 1990s and 2000s (see, for example, Bai et al., 2023). We thank Dan Lu for sharing the data with us. See Appendix A for more details about this data.

[20] We thank an anonymous referee for raising this point.

is indeed a great deal of heterogeneity in exporter behavior across sectors. The coefficient of variation of log exports ranges from 0.85 to 1.94, the export participation rate from 12 percent to 59 percent, the exit rate from 7 percent to 21 percent, and the incumbent premium from 1.76 to 4.82. Figure 6 plots these statistics against the mean sector-level NTR gap. There is a negative relationship between the NTR gap and the export participation rate, a positive relationship with the exit rate, and no clear relationship with either the coefficient of variation of log exports or the incumbent premium.

4.2 Assigned parameters and functional forms

Several of the model's parameters are assigned externally. A 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, $\tau_{gt}(1)$ and $\tau_{gt}(2)$, directly from the data described in Section 2.

The functional forms for the productivity process and death probability are taken from Alessandria et al. (2021). Firm productivity evolves according to

$$\log a' = \rho_z \ln a + \varepsilon, \qquad \varepsilon \overset{iid}{\sim} N(0, \sigma_{gz}^2), \tag{18}$$

where $z = \frac{1}{\theta_g - 1} \log a$. This specification eliminates the role of the demand elasticity, θ_g, in the size distribution of firms, which facilitates computation. We assume that the persistence parameter, ρ_z, is common across goods, while the variance of the innovations, σ_{gz}, varies.[21] The probability of death is $1 - \delta(a) = \max\left[0, \min\left(e^{-\delta_0 a} + \delta_1, 1\right)\right]$. We set ρ_z, δ_0, and δ_1 to the values reported in Alessandria et al. (2021). In the next section, we calibrate σ_{gz} to match the firm-size distribution in each sector.

We use the estimates of U.S. import demand elasticities in Soderbery (2018) to set the analogous elasticities in our model, θ_g.[22] We assume that θ_g is the same for all goods within

[21]The Chinese firm-level data cover too short a time period to accurately measure the persistence of idiosyncratic productivity shocks, let alone variation in this persistence across sectors. We therefore set this persistence externally. The variance of the innovations, however, is cleanly identified by cross-sectional variation in export sales.

[22]We use Soderbery (2018)'s estimates for U.S. imports at the product-origin level. The results are very similar when we use the product-level estimates for U.S. imports or the product-level estimates for all countries' imports.

a sector, but allow it to vary across sectors. We concord Soderbery (2018)'s estimates, which are reported at the HS4 level, to our 5-digit goods by taking the median within each good, and aggregating these good-level elasticities to our fifteen sectors by taking the average. Table 4 reports our sectoral demand elasticities, which range from 2.7 to 3.4. As Figure 6(e) shows, there is no systematic relationship between the demand elasticity and the NTR gap across sectors.

4.3 Calibrating the steady state to firm-level trade dynamics

We calibrate productivity dispersion, σ_{gz}, and non-tariff trade costs, f_{g0}, f_{g1}, ξ_{gH}, and ξ_{gL}, so that the model's terminal steady state matches the Chinese exporter-dynamics statistics computed in the 2004–2007 data (Table 3). Our empirical results show that the transition of Chinese exports to the United States was largely complete by 2004–2007, so we view these statistics as reflecting technological primitives rather than adjustments to policy changes (or changes in expectations about future policy). This allows us to calibrate the parameters that are identified by these statistics separately from the parameters that govern the transition.

We normalize the low variable trade cost, ξ_{gL}, to one in all sectors. There are a total of $15 \times 4 = 60$ parameters to calibrate at this stage, but the partial-equilibrium structure of our model allows us to calibrate each sector's parameters independently. These parameters are not individually identified, but each target statistic influences the identification of one parameter more than the others. Productivity dispersion, σ_{gz}, is primarily identified by the coefficient of variation of log exports. The entry cost, f_{g0}, is largely determined by the export participation rate. The continuation cost, f_{g1}, is mostly governed by the exit rate. The high variable export cost, ξ_{gH}, is pinned down by the incumbent premium. Table 4 lists the calibrated values of these parameters for each sector.

4.4 Calibrating the transition to aggregate trade dynamics

We calibrate the idiosyncratic probability of switching variable trade costs, ρ_ξ, and the probabilities of switching trade-policy regimes, $\{\omega_t(s, s')\}_{t=0}^{\infty}$, to match our empirical history of U.S.-China trade. Specifically, we target the long-run trade elasticity of -7.96 from the error-correction model and the annual NTR-gap elasticities shown in Figure 4.[23] To simulate the history of U.S.-China trade, we initialize our model so that all firms are non-nonexporters in 1970 (i.e., all firms have $\xi = \infty$), feed in the realized sequence of trade-policy regimes (NNTR from 1971 to 1979 and MFN from 1980 onward), update the distributions $\varphi_{g,t}$ using the laws of motion (14)–(16), and compute aggregate exports using (17). We then estimate equations (2) and (4) on the simulated data.

The parameters calibrated in this stage of our procedure are exactly identified: the number of parameters is the same as the number of target coefficients. The identification works as follows. The probability of switching trade costs, ρ_ξ, is primarily identified by the long-run aggregate trade elasticity.[24] As shown by Alessandria et al. (2021), this parameter governs the measure of high-capacity exporters (firms with $\xi = \xi_{gL}$) in the long run and thus plays a key role in determining the long-run response to trade reforms. We find $\rho_\xi = 0.91$.

The probability of switching from NNTR to MFN, $\omega_t(2,1)$, is identified by the NTR-gap elasticity during the 1970s, when China was in the NNTR regime. The intuition is that a higher probability of gaining MFN status boosts export participation more in high-gap industries than in low-gap industries because the former have more to gain from this status. Given the relatively small amount of time that China spent in the NNTR regime, coupled with the fact that the NTR-gap elasticity was relatively flat during this period, we assume for simplicity that $\omega_t(2,1)$ is constant and target the average NTR-gap elasticity between 1974–1979.

The probability of switching from MFN back to NNTR, $\omega_t(1,2)$, is identified by the NTR-gap elasticity from 1981 onward, which we HP-filter to smooth out temporary spikes (e.g., 1984). To be precise, the probability of losing MFN status in period t is identified by the NTR-gap elasticity in period $t+1$. As in Handley and Limão (2017), a higher probability of losing MFN status reduces export participation more in high-gap industries than in low-gap industries, but this does not affect trade volumes until the following period.[25]

[23] We HP-filter the NTR-gap coefficients from 1981 onward to smooth out temporary spikes (e.g., 1984). The calibrated model matches this smoothed series, shown in blue in Figure 7(a), exactly.

[24] Alternatively, we could target the long-run trade elasticity from our local-projections specification (3) instead of the ECM estimate. Because these elasticities are very similar, the implied parameter values would be very similar.

[25] To build intuition for the identification of $\omega_t(1,2)$, it is helpful to consider the version of the model in Appendix D, where firms always believe the current value of $\omega_t(1,2)$ will last forever, so that the yearly changes

It is important to emphasize that this stage of the calibration procedure is an indirect-inference exercise in the spirit of Gourieroux et al. (1993). As we discuss in Section 2, the reduced-form ECM and NTR-gap specifications, (2) and (4), are misspecified: the former is confounded by changes in expectations about future policy, while the latter is confounded by gradual adjustments to past changes in applied tariffs. These issues apply to the model as well as the data, and so the indirect inference approach allows us to use the estimates from these specifications to recover our model's structural parameters.

5 Results

We turn now to the results of our quantitative analysis. First, we discuss how trade-policy expectations have evolved since 1980. Second, we study the contribution of trade-policy uncertainty to the growth of Chinese exports to the United States. Third, we study the role of slow adjustment to the 1980 reform in explaining trade growth in subsequent decades. Fourth, we explore the importance of time variation in policy uncertainty and slow adjustments to earlier changes in expectations in explaining later patterns in trade growth. Last, we study the contribution of sectoral heterogeneity in exporter dynamics to our findings.

5.1 Estimates of trade-policy expectations

We begin with the main result of our calibration exercise: the annual probability of switching between policy regimes, which is shown in Figure 7(b). The probability of switching from the NNTR to the MFN regime is about 29 percent. The probability of switching back to the NNTR regime was initially 22.1 percent in 1980, rose sharply to 62.6 percent in 1981, before falling throughout the mid 1980s and early 1990s. There was a temporary increase during 1994–1996, followed by a smaller increase during the early 2000s, but the overall trend continued downward. By 2008, the end of our observation period, the probability of moving

in this probability are unanticipated. In this model, the probabilities from 1981 onward, $\{\omega_t(1,2)\}_{t=1981}^{2007}$, have no affect on the NTR-gap elasticity in 1981. This elasticity is entirely pinned down by $\omega_{1980}(1,2)$. Similarly, $\{\omega_t(1,2)\}_{t=1982}^{2007}$ do not affect the NTR-gap elasticity in 1981 or 1982. Since $\omega_{1980}(1,2)$ is already identified, we can use the 1982 elasticity to identify $\omega_{1981}(1,2)$. Iterating forward, we can uniquely identify all $\omega_t(1,2)$. In the baseline model, where firms know the entire path of probabilities, this one-for-one mapping between elasticities and probabilities breaks down somewhat, but both versions of the model yield very similar probability estimates. Thus, this mapping provides a good intuition for the identification in the baseline model.

back to the NNTR regime had fallen to 1.1 percent.[26]

Our estimates of trade-policy expectations in the years surrounding the 1980 reform are supported by the historical context.[27] Regarding the relatively low probability of gaining NTR status during the 1970s, although a few non-market economies had already gained NTR status, the Chinese case was complicated by the lack of formal diplomatic relations and political turnover in both the United States and China. The U.S. Presidency passed from Nixon to Ford and then to Carter during this period. In China, Hua Guofeng was appointed premier following the death of Zhou Enlai and Mao Zedong in 1976, and then Deng Xiaoping consolidated power in 1978. The United States established diplomatic relations with China in 1979, but there were still many additional steps to NTR involving Congress and the key question of Taiwan.

Regarding the likelihood of losing NTR status after 1980, the jump between 1980 and 1981 lines up with the change in U.S. leadership from Carter to Reagan; the latter was decidedly more hawkish on China than his predecessor, particularly concerning Taiwan and China's commitment to market reforms. Our finding of a substantial decrease in the probability of losing NTR status from 1985 to 1993 is consistent with several key policy reforms in China and changes in U.S.-China relations. In April 1984, Reagan visited China. In July of 1985, the United States and China signed an agreement on peaceful nuclear cooperation. In December 1985, the United States relaxed some export controls on technology. 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. As a further sign of an

[26] That the probability does not fall to zero in the long run implies that trade is permanently depressed, albeit slightly, by the possibility of losing NTR status. This is consistent with recent evidence from the U.S.-China trade war, which shows that trade in goods with high NTR gaps rose relative to trade in low-gap goods when President Trump introduced a new set of tariffs on Chinese goods in late 2018. These tariffs were essentially orthogonal to the NTR gap, so the change in the NTR-gap elasticity during this period suggests a reduction in the probability of losing NTR status. One interpretation of this evidence is that the U.S.-China trade relationship had shifted into a new paradigm where the uncertainty was no longer about losing NTR status, but about whether—and how long—the trade war would last. We study policy uncertainty during the trade war in Alessandria et al. (2023b).

[27] Much of this discussion is based on the series of annual reports that provide summaries of key changes in trade and trade policy (USITC, 1970–1990). We have also drawn from special reports on China and the U.S.S.R. (USITC, 1977a; 1977b); a report on China's development and its impact on the U.S. economy (USITC, 1985); a series of reports published by the U.S. Department of Commerce (U.S. Department of Commerce, 1977–1985); a series of declassified studies on potential trade with China going back to the 1950s (CIA, 1950–1980); and the White House Historian's "History of Foreign Relations."

improving U.S.-China relationship, Deng Xiaoping was named Time's Man of the Year for the second time in 1985.

One of our most notable findings is the lack of a discrete drop or accelerated decline in the probability of moving back to the NNTR regime when China gained PNTR status in 2001. This probability was about the same in the mid-2000s as in the early 1990s, and rose temporarily by several percentage points during 2000–2003. This finding tells a different story about the effect of gaining PNTR status than Pierce and Schott (2016) and Handley and Limão (2017), who argue that the change in the NTR-gap elasticity after 2001 is evidence that PNTR significantly reduced the probability of losing NTR status. There are two key factors in explaining this difference. First, in Section 5.2, we show that gradual adjustment to the 1980 NTR grant (and the 1971 lifting of the embargo), plays an important role. Even in the absence of policy uncertainty, the NTR-gap elasticity would still have fallen after 2001. Second, in Section 5.4, we show that part of the decline in the NTR-gap elasticity after 2001 was driven by a gradual adjustment to the earlier changes in policy uncertainty that occurred in the late 1980s and early 1990s. When we do not allow for these earlier changes, our model requires a larger reduction in the probability of losing NTR status in 2001.

One way to understand our trade-policy probability estimates is to compare the realized path of tariffs with the mean discounted expected tariff that Chinese exporters faced at each point in time. Figure 7(b) plots the discounted expected tariff across goods in the model,

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

alongside the mean applied tariff. Whereas the realized path of applied tariffs fell sharply in 1980, and then fell slightly throughout the 1980s and 1990s due to continued reforms to U.S. MFN tariff rates, the discounted expected tariff fell gradually, but steadily, throughout the entire period. The discounted expected tariff remained above the applied MFN rate even after China joined the WTO in 2001.

The dynamics of the discounted mean expected tariff help us understand the gradual adjustment of export volumes to the abrupt decrease in the current tariff. Recall that the intensive margin of trade, exports per exporter, is mostly determined by the current tariff and the distribution of variable trade costs, whereas the extensive margin, through entry and exit decisions, is determined by the path of future tariffs. The slower to decrease discounted expected tariff suppressed the participation of Chinese firms in the export market. As the discounted mean expected tariff fell, export participation increased and aggregate trade volumes grew.

5.2 The effects of policy uncertainty on trade

To measure how changes in policy expectations have affected trade flows, we compare our benchmark model with a counterfactual one in which firms believe the current trade-policy regime is permanent (equivalent to setting $\omega(2,1)$ and $\omega_t(1,2)$ to zero). In this *no-TPU counterfactual*, firms believe during 1971–1979 the NNTR regime will last forever. In 1980, firms are surprised by the shift to the MFN regime, but from then onward they believe the MFN regime will last forever.[28] 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.[29]

Figure 7(a) plots the NTR-gap elasticities obtained by estimating (4) on simulated data from the no-TPU counterfactual. During the 1970s, the elasticities in the counterfactual are lower than in the benchmark, where the possibility of gaining MFN status boosted 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. The difference between the counterfactual elasticity and the actual elasticity at each point in time measures the contribution of policy uncertainty. This contribution was about 50–70 percent of the elasticity in the data during the 1980s, 50–60 percent during the 1990s, and 65–80 percent during the 2000s. Note that policy uncertainty actually accounted for a smaller portion of the NTR-gap elasticity during the periods when NTR reversal was most likely (the late 1980s), and a larger portion of the gap when reversal was least likely (the 2000s). This is because delayed adjustment to the 1980 reform played

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

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

a more important role in the earlier period, when the liberalization was newer.

The no-TPU counterfactual also allows us to assess the impact of trade-policy uncertainty on aggregate trade. Figure 7(d) plots aggregate exports in the counterfactual and the benchmark models. Again, the vertical distance between the two lines measures the effect of trade-policy uncertainty. The results largely mirror those in Figure 7(a) with one notable exception: a material difference remains in aggregate trade between the counterfactual and the benchmark after China joins the WTO in 2001. The reason is that the probability of losing NTR status is positive (albeit small) even in the long run. The long-run discounted expected tariff in the benchmark model is higher than the applied tariff (Figure 7(c)), which permanently reduces the number of exporters in high-gap industries. This is consistent with our observation in Alessandria et al. (2023b) that exports of high-gap goods grew relative to exports of low-gap goods after the onset of the U.S.-China trade war in 2018, even though the "trade-war gap" (the difference between the trade-war tariffs and MFN tariffs) is orthogonal to the NTR gap.

5.3 The role of slow adjustment

The no-TPU counterfactual also allows us to measure the role of slow adjustment to the 1980 NTR grant (and the 1971 lifting of the embargo) in explaining trade growth, particularly the patterns in the 1990s and 2000s that are commonly attributed to a reduction in trade-policy uncertainty caused by PNTR access.

To do so, we estimate the pre-PNTR elasticity of trade to the NTR gap using the same specification and 1992–2007 observation period as Pierce and Schott (2016):

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

When we estimate this regression using the observed data, we find $\beta = -0.92$, and we obtain essentially the same estimate using simulated data from our benchmark model with policy uncertainty.[30] When we estimate this regression on simulated data from the no-TPU counterfactual, we find $\beta = -0.26$, which is smaller than the estimate from the data but

[30] Our estimate is different than Pierce and Schott (2016)'s because we use a different level of aggregation. See Appendix C for additional details.

still statistically significant. This shows that even if there had never been doubt about the permanence of China's NTR status, exports of high-gap goods still would have grown faster than imports of low-gap goods after China gained PNTR status in 2001. Gradual adjustment to the earlier reforms accounts for more than one-quarter of the overall effect of PNTR on trade documented by Pierce and Schott (2016).

A different way to study the role of gradual adjustment is to examine how our estimates of trade-policy transition probabilities differ in an alternative model, with no exporter life cycle, in which trade adjusts more quickly. In this *fast-adjustment model*, we replace 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. In this model, which is similar to the model used by Handley and Limão (2017), we find much larger estimates of the non-renewal probability and a much smaller likelihood of transitioning from NNTR to MFN.

In Figure 8(a), we plot the elasticity of trade to the NTR gap in the benchmark and fast-adjustment model, with and without TPU. The elasticities in the fast-adjustment model without TPU converge to zero faster than in the benchmark model without TPU. This implies that, in the fast-adjustment model with TPU, higher probabilities of switching back to the NNTR regime are required to match the observed elasticities—in fact, the simple dynamic model requires this probability to be one in 1980–1981 (Figure 8(b)). This is because the fast-adjustment model places less weight on slow adjustments, and thus more weight on policy expectations, in accounting for the dynamics of U.S. imports from China.

5.4 The role of time-varying policy uncertainty

Our quantitative strategy leverages year-by-year variation in the elasticity of trade to the NTR gap to identify changes in expectations about U.S. trade policy toward China, and we find that most of these changes occurred in the late 1980s and early 1990s, rather than, as others have argued, at the turn of the century. Here, we ask, What would happen if we assumed only one or two changes in expectations occurred 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 before falling to zero in 2001.[31] We calibrate the probability of losing NTR status during 1980–2000 to match the single NTR-gap elasticity estimated by Pierce and Schott (2016), over 1992–2007, from (20). Figure 9(a) shows this version of the model (*Const. TPU from 1980*) does fairly well in matching the evolution of the NTR-gap elasticity during the 1990s and 2000s. Trade grows too quickly, however, in high-gap industries relative to low-gap industries in the 1980s. As Figure 9(b) shows, the probability of losing NTR status in this version of the model is about 5 percent, which is more than twice the decline in the average probability in the benchmark model between 1992–2000 and 2001–2007.

We also consider a version of our model 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 1990, which is a common assumption in the trade-policy uncertainty literature.[32] Figure 9(a) shows this model (*Const. TPU from 1990*) also does a good job of matching the trajectory of the NTR-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 NTR-gap elasticity in this model is clearly at odds with the data. The probability of losing NTR status in this model is almost 20 percent, which is well above the highest probability in the 1990s benchmark-model estimates and more than ten times the difference between the 1992–2000 and 2001–2007 averages.

These analyses show that 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. Rising credibility of U.S. trade policy toward China during the mid-to-late 1980s was an important factor in explaining the growth of U.S. imports

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

[32] For example, according to Bianconi et al. (2021), "[a]nnual 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).

from China over the next two decades, and ignoring this trend leads one to overstate the degree to which uncertainty fell after China gained PNTR status in 2001. They also highlight an important lesson that has broad applicability outside of the U.S.-China context: trade adjusts slowly to changes in expectations about future policy as well as past changes in tariffs.

5.5 The role of sectoral heterogeneity

In our calibration, we grouped goods into fifteen sectors and allowed the parameters that govern the dynamics of exporting firms to vary across sectors. There are no clear relationships between these parameters and the NTR gap, which indicates that this sectoral heterogeneity is unlikely to play a large role in our results. To demonstrate this, we compare our benchmark model to an alternative model in which we turn off this sectoral heterogeneity. In the *one-sector model*, demand elasticities, non-tariff trade costs, and productivity dispersion are the same for all goods. We calibrate these parameters to match aggregate exporter-dynamics statistics from China (i.e., we compute these statistics for the entire dataset, rather than sector by sector).

Figure 10 compares the results in the one-sector model to the benchmark results. The probability of gaining NTR status during the 1970s in the one-sector model is about 12 percent, less than half of the probability in the benchmark model. In the absence of policy uncertainty, the NTR-gap elasticity during the 1970s is smaller—closer to the actual data—in the one-sector model than in the benchmark. This means that less anticipation of gaining NTR status is needed to fit the data during this period. Conversely, the probability of losing NTR status from 1980 onward is higher in the one-sector model than in the benchmark during the 1980s, but there is no material difference from the late 1980s onward.

Overall, these results show that sectoral heterogeneity in the technological primitives that determine exporter behavior does not play a large role in driving our results. Ignoring this heterogeneity would lead us to overstate the likelihood of losing NTR status in the early 1980s, but would not affect our main findings that this likelihood fell dramatically during the late 1980s and changed little when China gained PNTR status in 2001.

6 Employment effects

Our estimates imply that changes in policy uncertainty played the largest role in export growth during 1986–1993, and the effect in this period was much larger than the effect from China joining the WTO in 2001. At first glance, this narrative appears inconsistent with the findings of Pierce and Schott (2016), who document a large decline in U.S. employment in high-gap industries following China's accession to the WTO. In this section, we revisit these employment effects by studying the elasticity of U.S. employment to the NTR gap over our long sample period. Once we introduce the industry controls suggested by theory, the decline in employment related to trade policy on Chinese imports starts much earlier and does not accelerate when China joins the WTO. Instead, most of the drop in employment in the period around China's WTO accession appears to be related to industry-specific factors that are correlated with the NTR gap, rather than import growth caused by a change in the credibility of U.S. trade policy towards China.

6.1 Conceptual framework

To set ideas, consider an Armington model of U.S. demand with an aggregate consumption good, Q_t, that is a bundle of industry-level goods, Q_{gt}, with elasticity of substitution α,

$$Q_t = \left(\sum_g Q_{gt}^{\frac{\alpha-1}{\alpha}} \right)^{\frac{\alpha}{\alpha-1}}. \tag{21}$$

Each good g is a combination of U.S. domestically-produced goods, Q_{Dgt}, and Chinese imports, Q_{Mgt}, with elasticity of substitution θ_g,[33]

$$Q_{gt} = \left(Q_{Dgt}^{\frac{\theta_g-1}{\theta_g}} + Q_{Mgt}^{\frac{\theta_g-1}{\theta_g}} \right)^{\frac{\theta_g}{\theta_g-1}}. \tag{22}$$

[33] This structure is consistent with the demand function in (6).

The demand function for the domestic good and the aggregate price of good g are

$$P_{Dgt}Q_{Dgt} = \left(\frac{P_{Dgt}}{P_{gt}}\right)^{1-\theta_g} P_{gt}Q_{gt}, \tag{23}$$

$$P_{gt} = \left(P_{Dgt}^{1-\theta_g} + (P_{Mgt}\tau_{Mgt})^{1-\theta_g}\right)^{\frac{1}{1-\theta_g}}, \tag{24}$$

where P_{Dgt} is the price of the domestically-produced good g, P_{Mgt} is the price of Chinese exports, τ_{Mgt} is the tariff levied on Chinese exports, and P_{gt} is the price of good g. The price of Q_t is normalized to one. Sales by domestic producers is the sum of domestic and export sales, $P_{Dgt}Q^*_{Mgt}$,

$$P_{Dgt}Y_{gt} = P_{Dgt}Q_{Dgt} + P_{Dgt}Q^*_{Mgt}, \tag{25}$$

where Q^*_{Mgt} is the quantity demanded in China and we have assumed the export price and domestic price are identical.

After combining (23), (24), and (25), and log-linearizing, we arrive at the relationship between real good-level sales by U.S. producers and good-level trade policy,

$$\begin{aligned} d\ln Y_{gt} \approx &(1-\omega_{Xg})[\omega_{Mg}(\theta_g - 1)(d\ln(P_{Mgt}\tau_{Mgt}) - d\ln P_{Dgt}) \\ &+ d\ln(P_{gt}Q_{gt}) - d\ln P_{Dgt}] + \omega_{Xg}\left[d\ln(P^*_{Mgt}Q^*_{Mgt}) - d\ln P_{Dgt}\right], \end{aligned} \tag{26}$$

where ω_{Xg} is industry g's export share of total sales and ω_{Mg} is the Chinese import share of total domestic absorption, both of which are computed in the linearization's base period. Assuming production in U.S. industry g is $Y_{gt} = Z_{gt}L_{gt}$, where Z_{gt} is industry g's labor productivity, (26) becomes

$$\begin{aligned} d\ln L_{gt} \approx &(1-\omega_{Xg})[\omega_{Mg}(\theta_g - 1)(d\ln P_{Mgt}\tau_{Mgt} - d\ln P_{Dgt}) \\ &+ d\ln(P_{gt}Q_{gt}) - d\ln P_{Dgt}] + \omega_{Xg}\left[d\ln(P^*_{Mgt}Q^*_{Mgt}) - d\ln P_{Dgt}\right] - d\ln Z_{gt}. \end{aligned} \tag{27}$$

This equation shows that the change in employment in industry g depends on trade policy, through its effect on relative import prices, domestic absorption, foreign sales, and labor productivity.

Notice that (27) implies the employment effects of trade policy are amplified or attenuated by the industry's exposure to imports and exports, as captured by ω_{Xg} and ω_{Mg}. The same change in Chinese imports due to a trade liberalization will have a larger effect on employment in an industry in which imports initially make up a larger share of domestic absorption, and a larger effect in an industry that sells most of its output domestically. This heterogeneous effect of trade policy and import substitution is key to reconciling our results on trade with previous findings in the literature on employment.

6.2 Estimation details

Relative to our work with U.S. import data, there are two challenges in estimating (27). First, our data on U.S. sales and employment, which come from Becker et al. (2021), are measured at a higher level of aggregation, the Standard Industrial Classification (SIC).[34] One advantage of using these more aggregated data, however, is their availability from 1958 onward, which allows us to study employment trends for more than a decade before the United States opened to trade with China, when it is unlikely that the prospect of import growth in the far future was a material factor.

The second concern is the lack of data on the price of Chinese imports relative to domestically-produced goods, $d\ln P_{Mjt}\tau_{Mjt} - d\ln P_{Djt}$. The literature, however, focuses primarily on the change in employment from import growth caused by changes in expectations about future trade policy, which is commonly understood to be captured by the relationship between employment and the NTR gap. Recognizing that this relationship also captures the effects of import growth resulting from gradual adjustment to the 1980 reform, we replace $d\ln P_{Mjt}\tau_{Mjt} - d\ln P_{Djt}$ by the NTR gap interacted with an annual indicator variable as in (4):

$$
\begin{aligned}
d\ln L_{gt} = &(1-\omega_{Xg})\left[\omega_{Mg}\sum_{t'=1958}^{2007}\beta_t \mathbb{1}_{\{t=t'\}} GAP_g + d\ln(P_{gt}Q_{gt}) - d\ln P_{Dgt}\right] \\
&+ \omega_{Xg}(d\ln(P^*_{Mgt}Q^*_{Mgt}) - d\ln P_{Dgt}) - d\ln Z_{gt} + \delta_g + \delta_t + \epsilon_{gt},
\end{aligned} \tag{28}
$$

[34]Reduced-form estimates are nearly identical using the North American Industrial Classification System.

where we have further included industry (δ_g) and time (δ_t) fixed effects.[35]

Notice that the shares (ω_{Xg}, ω_{Mg}) are time invariant. As discussed above, these shares are measured in the base period around which we linearize. Figure 11 plots the distributions of the industry shares over time. Figure 11(a) shows that the import shares from China were skewed towards zero in the earlier period and grew over time. This is consistent with the Chinese import penetration documented in Section 2. Over the same period, the domestic sales share of U.S. producers fell and the distribution shifted left, as seen in Figure 11(b). We use the average shares over 1995–1999 for ω_{Xg} and ω_{Mg}, which serve as the linearization base period.[36] We compute these shares by aggregating the SITC trade data to the SIC level, which is straightforward.[37]

6.3 Estimation results

Our baseline specification and several alternatives are plotted in Figure 12. The industry-specific shares in (28) imply that we estimate one set of parameters for each industry in our data. In Figure 12, we report results for the industry in the 90th percentile of Chinese import share and the 50th percentile export share.[38] Plotted in Figure 12 is $\beta_t(1-\omega_{Xg})\omega_{Mg}$, which is the elasticity of employment to the NTR gap. We find that employment falls relatively more in industries that are more exposed to the NTR gap, but that most of the decline occurs before China joins the WTO.

How do our findings compare with those in Pierce and Schott (2016)? Following their approach, we estimate the following reduced-form equation,

$$d \ln L_{gt} = \sum_{t'=1958}^{2007} \beta_t \mathbb{1}_{\{t=t'\}} GAP_g + \delta_g + \delta_t + \epsilon_{gt}. \tag{29}$$

[35] By imposing the same elasticity on all industries, this specification ignores heterogeneity in θ_g. However, θ_g is approximately orthogonal to the NTR gap, so this heterogeneity should not affect our estimate of β_t (Figure 6(e)).

[36] Our results are robust to using average shares over 1975–1979 or the average across the full sample for which trade data is available at the country-good-year level, 1974–2008.

[37] SIC-level trade data is available from 1958 onward. Data on applied tariffs at the SIC level are incomplete until 1974, but we do not use the tariff data in this part of the paper.

[38] In the appendix, Figure G.12 plots the dynamics of employment for different Chinese import penetration and U.S. producer domestic market shares. Given the large variation in import penetration, we find substantial variation in the employment effects as we change ω_M. Given the much smaller variation in domestic sales shares, we find less variation in the employment effects when we change ω_X.

Relative to (28), this specification forces the coefficients employment elasticity to be constant across industries (i.e., forces ω_{Xg} and ω_{Mg} to be the same in all industries) and drops the industry controls for domestic expenditures, exports, and TFP. Figure 12 shows that without these controls, we estimate very different dynamics: Employment is stable during 1990 to 1997, contracts gradually through 2001, and then declines sharply. We also find robust growth in employment that is correlated with the NTR gap from 1958–1980, which is counterintuitive, given that China was a small part of the U.S. economy in this period. The large differences in moving from the theoretical regression in (28) to the reduced-form regression in (29) suggests that the decline in employment in high-gap industries around the time that China joined the WTO may be spurious.

We gain some insight into the importance of the theory-motivated controls by considering alternative versions of (28) in Figure 12. Removing TFP changes our results little. Removing TFP and the controls for export and domestic demand leads to a slightly smaller long-run effect on employment, but by 1994, the effects are almost identical to the baseline model. Dropping the industry shares as well as TFP, export and domestic demand, leaves us with (29), which is very different from the baseline. Thus, heterogeneity across industries in import and export exposure, and the way this heterogeneity modulates the effect of import growth on employment, plays the key role in reconciling our findings with those of Pierce and Schott (2016).

6.4 Understanding the results

Why does removing the theory-motivated controls change the employment elasticities so dramatically? The difference between our specification and that in (29) can be further understood by studying the components of the employment elasticity, $\beta_t(1-\omega_{Xg})\omega_{Mg}$.

Industry shares. Our results in Figure 12 change the most when we do not control for heterogeneous import and export exposure through ω_{Xg} and ω_{Mg}. Figure 13 plots the distribution of NTR gaps by quartile of average Chinese import share and domestic sales share during 1995–1999. There is substantial variation in the spreads across the quartiles. Importantly, industries with higher Chinese import shares tend to have larger and less dispersed gaps. We would have expected the opposite: industries facing large NTR gaps should have smaller import shares as TPU suppresses entry by Chinese firms. This suggests that these industries with large NTR gaps and large import shares are industries in which China has a comparative advantage. Thus, it is important to allow the elasticity of employment to the NTR gap to vary with import and export exposure, lest we attribute employment declines in high gap industries to multilateral China supply factors.

Substitution elasticities. The NTR-gap elasticities, β_t, that we estimate with the employment data should, in theory, be the same as the elasticities (with the opposite sign) that we estimate with the import data from (4). Absent issues of aggregation, the main differences are that (28) lacks many of the rich controls available with the import data and the substitution identified is between imports and domestic sales, rather than between imports from different source countries.

We compare our estimates of β_t from the trade and employment data in Figure 14. Using employment data, we estimate NTR-gap elasticities that are about twice as large as those from the import data. Moreover, the employment data suggests that the substitution comes much later than in the trade data. To reconcile these surprising findings, we remove several of the fixed effects from our trade regression that are omitted from our employment regression. We estimate (4) with only time fixed effects and product-country fixed effects. With these minimal fixed effects, we find import substitution towards high-gap goods that is almost twice as large as our baseline trade estimate, and coincides, both quantitatively and temporally, with our estimates from the employment data. That the estimates from the trade data can be made consistent with the estimates from the employment data by omitting standard fixed effects points to some concerns with using employment to identify the effects of Chinese trade policy.

Since the gap elasticities from the trade estimation are measuring the same substitution forces as in the employment regressions, we can use them to compute an alternative measure of the effect of the NTR gap on employment. In Figure 12, we plot $\beta_t(1-\omega_{Xg})\omega_{Mg}$ where we use the estimates of β_t from the import regression (4) ("using trade coefficients" in the figure) instead of the estimates from the employment regression (28). The effects on employment are half as big as those in our baseline results using employment data.

Industry trends. We now take a broader view of trends in variables that are correlated with the NTR gap that could influence both our theoretically-consistent regression (28) and our reduced-form specification (29). Figure 15 plots the elasticity of several key variables to the NTR gap from 1958–2018. Figure 15(a) plots the dynamics of exports, imports, and the applied tariff; note that we are plotting total industry-level imports, not just imports from China. The applied tariff series, which starts in 1974, falls with the NTR gap as high-gap industries also had relatively large tariff declines. These differential reforms seem to end around 2001. U.S. imports and exports rose more in high-gap industries, and this growth is particularly pronounced in imports. Imports grew faster from 1982 to 2001 but, since then, it has mostly been in line with other goods. The faster growth in imports is to be expected given the tariff declines.

Figure 15(b) plots the dynamics of employment and domestic absorption. These series generally move together, with employment and domestic absorption growing robustly in the early years of the sample. Since the late 1990s, employment and domestic absorption have fallen sharply. The relative growth in high-gap industries in the early part of the sample is about half as large as the decline at the end of the sample. Obviously, the growth in the early part of the sample cannot be attributed to China, since China was still under embargo. Moreover, the dynamics of these series since the late 1990s—modest movements in trade coupled with the large movements in domestic absorption—suggest that there may be other industry-level shocks that are more important in accounting for differences in employment across sectors. Without a multisector general-equilibrium model, however, which is beyond the scope of this paper, it is not possible to say more.

7 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 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 PNTR status in 2001. This observation suggests much of the growth in trade in products with high NTR gaps in the 1990s and 2000s was a delayed effect of earlier liberalizations—and earlier increases in their perceived credibility. It also indicates that the initial lack of credibility about the 1980 NTR grant depressed trade much more than later concerns about NTR status renewal that were eliminated when China gained PNTR status in 2001.

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, whereas 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 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 toward 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, because these key aspects of U.S. policy toward 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. (2023a) 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		2008	
	Sector	Mean	Std.	1979	2008	China	NTR	China	NTR
1	Food, beverage and tobacco	24	20	3	1	23	6	4	3
2	Textile, clothing, leather and footwear manufacturing	25	16	10	5	32	10	4	5
3	Wood and straw products	28	13	0	1	33	5	2	1
4	Paper and printing products	20	13	0	2	18	3	0	0
5	Energy products and chemicals	30	17	20	4	18	4	2	2
6	Rubber and plastic products	61	19	0	4	71	5	3	2
7	Non-metallic mineral products	44	17	1	1	45	10	3	2
8	Base metal manufacturing	24	20	2	3	20	3	1	1
9	Calendered metal manufacturing	44	17	3	9	42	6	3	2
10	Other machinery and equipment manufacturing industry	32	14	0	24	30	5	1	1
11	Computer, electronic and optical products	55	29	0	2	40	9	2	2
12	Electrical equipment manufacturing	35	5	0	23	35	6	2	1
13	Vehicle manufacturing	30	0	0	1	40	4	1	1
14	Furniture and other manufacturing	47	19	6	14	49	7	3	2
15	Non-manufacturing	10	19	54	6	17	5	1	1
	Average	29	22	16	8	28	6	2	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 an FTA with it (see footnote 6). When computing the mean applied duties of NTR countries, only SITC codes with nonzero U.S. imports from China are considered.

Table 2: Calibration summary

Parameter	Meaning	Value	Source/target
(a) Assigned			
r	Interest rate	4 pct.	Standard
ρ_z	Persistence of productivity	0.65	Alessandria et al. (2021)
δ_0	Corr.(survival,productivity)	21.04	Alessandria et al. (2021)
δ_1	Minimum death probability	0.023	Alessandria et al. (2021)
τ_{g1}	NNTR tariff	Varies	Data
τ_{g2}	MFN tariff	Varies	Data
θ_g	Demand elasticity	Varies	Soderbery (2018)
(b) Calibrated in steady state			
f_{g0}	Entry cost	Varies	Export participation rate
f_{g1}	Continuation cost	Varies	Exit rate
ξ_g	High iceberg cost	Varies	Incumbent premium
σ_{gz}	Productivity dispersion	CV of log sales	
(c) Calibrated to transition			
ρ_ξ	Prob. of keeping iceberg cost	0.87	ECM estimate of LR trade elasticity = -7.96
$\omega(2,1)$	Prob. NNTR to MFN	0.29	Avg. NTR-gap elasticity during 1974–1979
$\omega_t(1,2)$	Prob. MFN to NNTR	Varies	NTR-gap elasticity during 1981–2008

Table 3: Chinese exporter dynamics statistics, 2004–2007

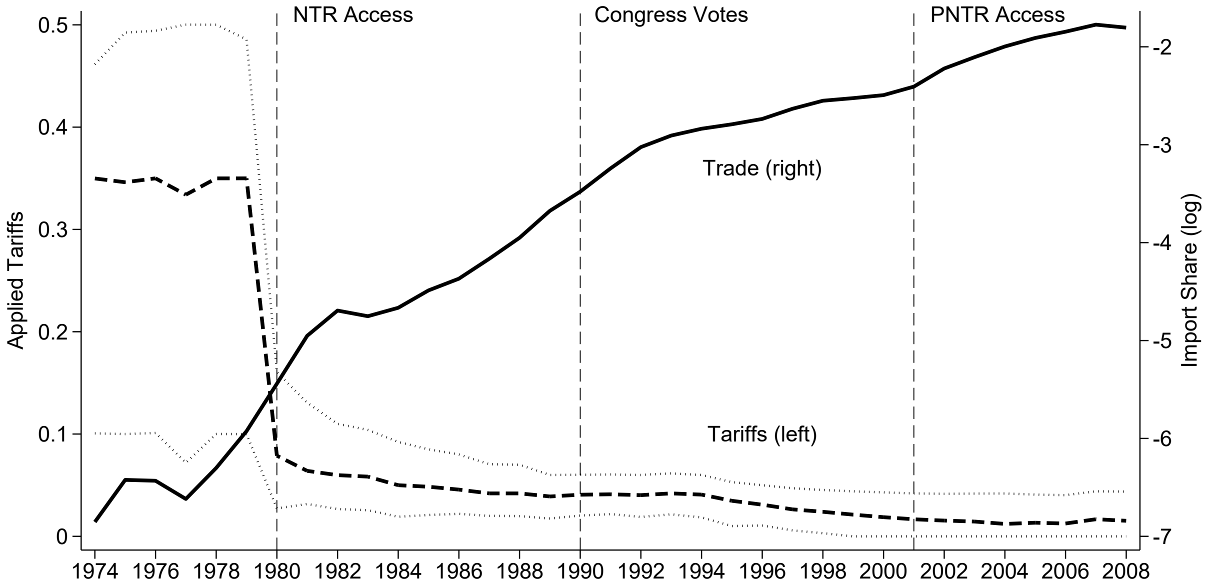
	Sector	Export part.	Exit rate	Incumbent size prem.	Log CV exports
1	Food, beverage and tobacco	19	16	2.71	0.91
2	Textile, clothing, leather and footwear manufacturing	45	10	1.99	1.06
3	Wood and straw products	24	13	2.05	1.09
4	Paper and printing products	12	17	3.10	1.30
5	Energy products and chemicals	19	15	3.23	1.48
6	Rubber and plastic products	29	10	2.69	1.08
7	Non-metallic mineral products	16	18	2.26	0.85
8	Base metal manufacturing	12	21	3.96	1.15
9	Calendered metal manufacturing	29	10	2.48	1.24
10	Other machinery and equipment manufacturing industry	23	13	3.33	1.54
11	Computer, electronic and optical products	48	7	4.82	1.94
12	Electrical equipment manufacturing	32	10	3.35	1.55
13	Vehicle manufacturing	23	12	4.07	1.31
14	Furniture and other manufacturing	59	7	1.76	0.95
15	Non-manufacturing	28	13	2.99	1.25

Notes: The moments reported here are obtained using firm-level data from Chinese manufacturers. The data is described in Appendix A. All moments refer to sectoral averages between 2004 and 2007. Export participation, in percent, is calculated as the number of firms with positive export sales over the total number of firms in a sector. The exit rate, in percent, is calculated as the number that exported in $t-1$ but not in t over the number of exporters in t. The incumbent size premium is calculated as the sales of incumbent firms over sales of new exporters. The last column is the log value of the coefficient of variation of sectoral export revenues.

Table 4: Sector-level model parameters

	Sector	θ_g	f_{g0}	f_{g1}	ξ_{gH}	σ_{gz}
1	Food, beverage and tobacco	3.09	0.13	0.12	3.30	0.91
2	Textile, clothing, leather and footwear manufacturing	3.02	0.20	0.13	2.59	1.02
3	Wood and straw products	2.78	0.26	0.17	3.71	1.03
4	Paper and printing products	3.44	0.19	0.16	3.47	1.08
5	Energy products and chemicals	2.99	0.27	0.20	4.56	1.17
6	Rubber and plastic products	3.13	0.19	0.12	3.40	0.99
7	Non-metallic mineral products	2.84	0.16	0.14	3.55	0.90
8	Base metal manufacturing	3.04	0.13	0.16	4.59	0.99
9	Calendered metal manufacturing	2.73	0.30	0.17	4.62	1.08
10	Other machinery and equipment manufacturing industry	3.74	0.23	0.16	3.04	1.20
11	Computer, electronic and optical products	3.16	0.46	0.21	4.84	1.36
12	Electrical equipment manufacturing	3.27	0.31	0.16	3.88	1.20
13	Vehicle manufacturing	3.14	0.21	0.16	4.92	1.07
14	Furniture and other manufacturing	3.22	0.20	0.11	2.26	0.98
15	Non-manufacturing	2.97	0.22	0.16	4.04	1.07

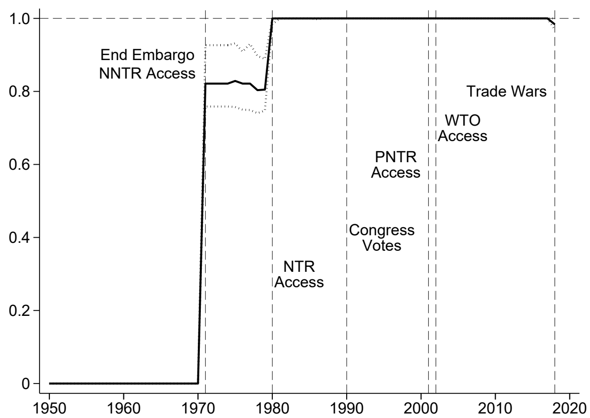
Figure 1: China's aggregate import share and duties



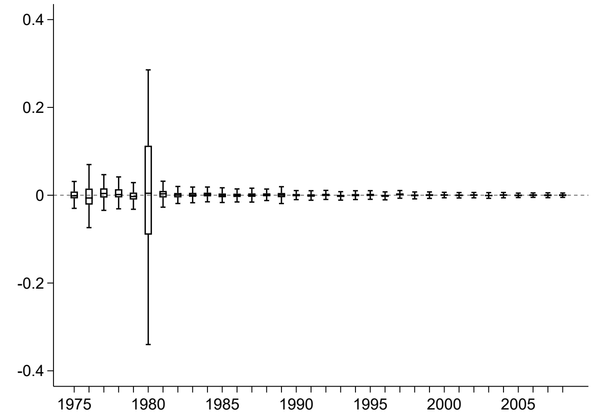
Notes: The dashed line is the median tariff across goods and the dotted lines are the 25^{th} and 75^{th} percentiles. The solid line is the log import share of China in total U.S. imports.

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

(a) Inverse tariff



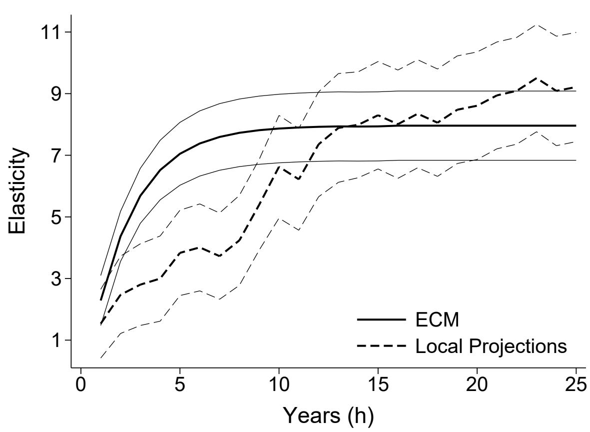
(b) Residual variation of tariff changes



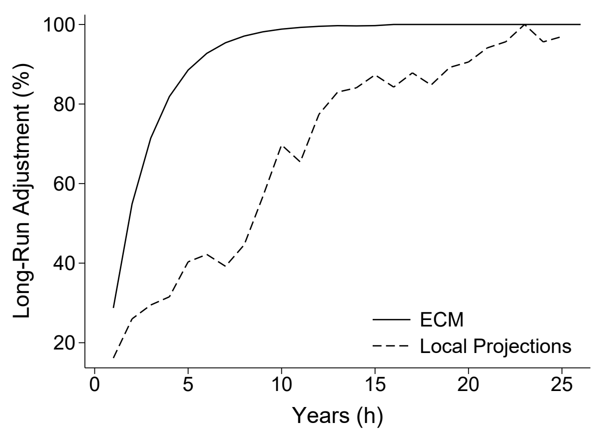
Notes: Panel (a) plots China's inverse tariff defined in (1). The solid line is the median of all products and the dashed lines are the 25th and 75th percentiles. Panel (b) plots the distribution of the annual residual variation in China's tariff changes by considering the residual from regressing $\Delta_1 \tau_{jgt}$ 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

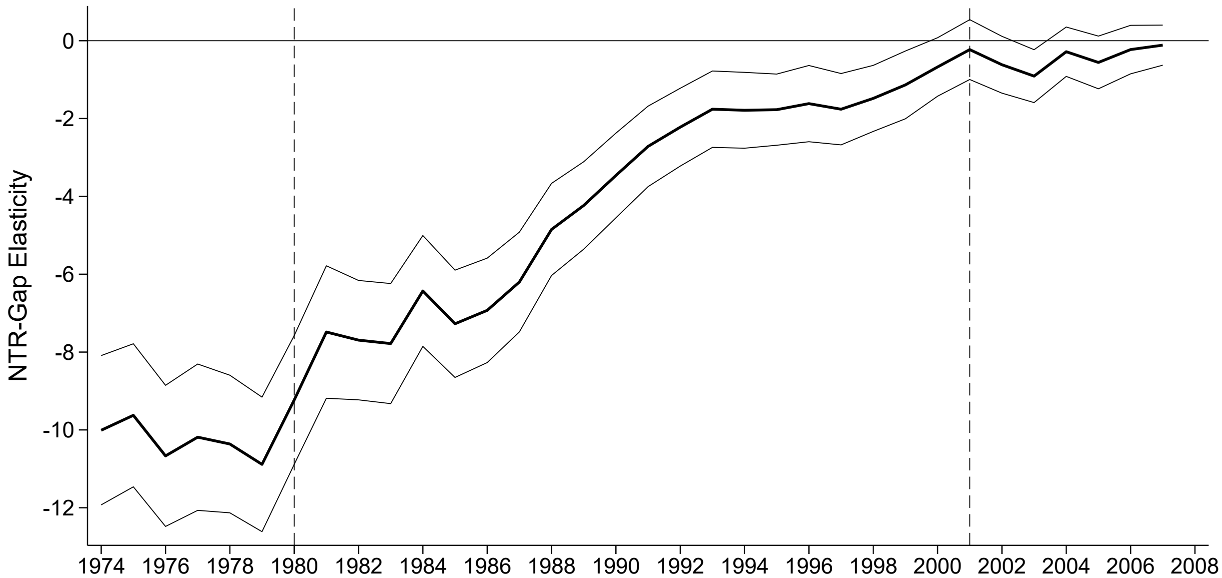


(b) Speed of adjustment



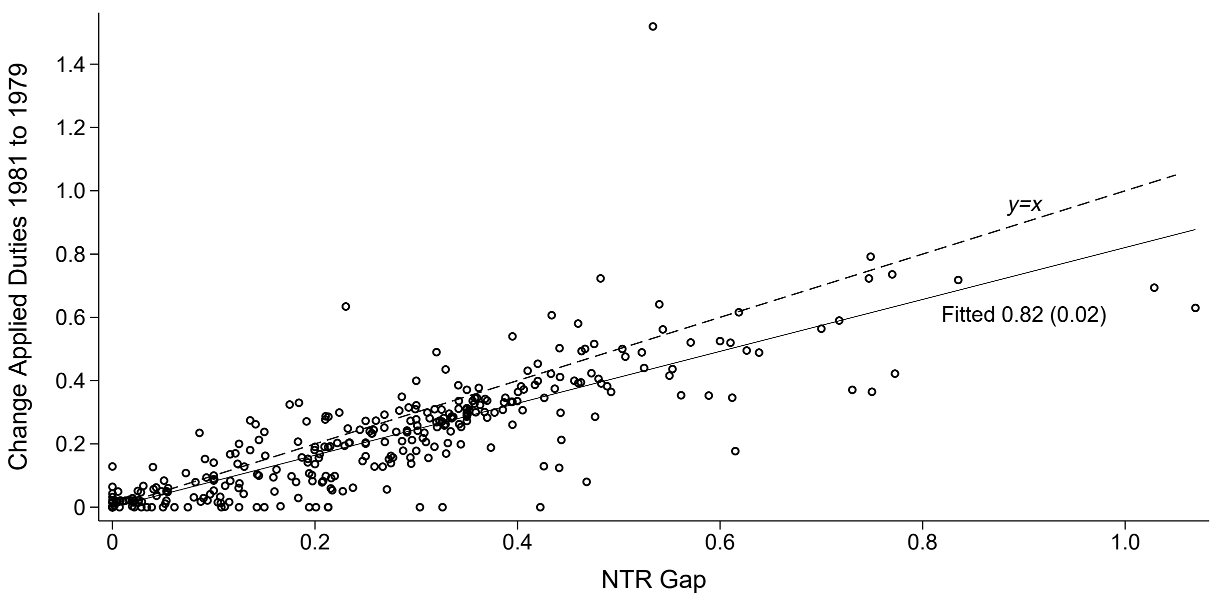
Notes: Panel (a) reports the time-varying elasticities to tariff changes on U.S. imports from China. The solid line uses the ECM approach of (2) and the dashed line uses the local-projections approach of (3). The standard errors that construct the 95-percent confidence intervals are clustered at the 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 the NTR gap



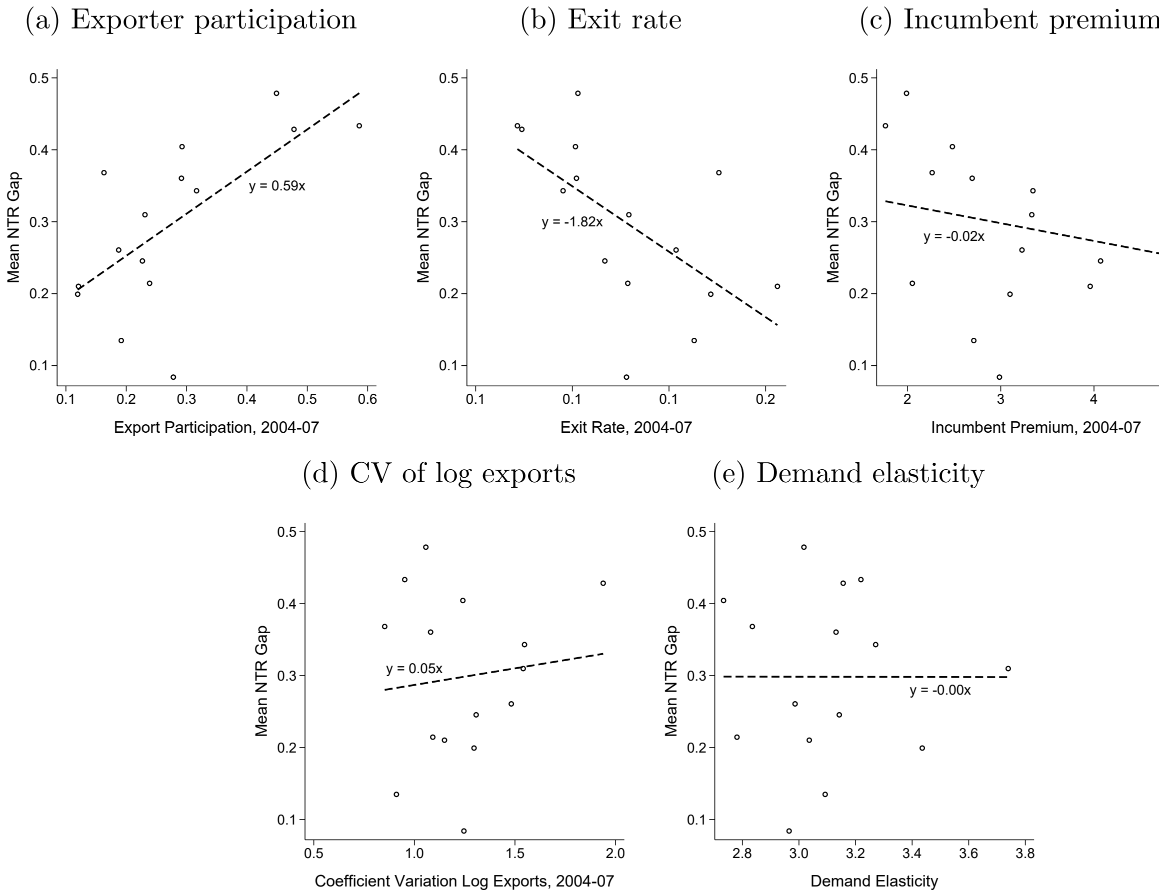
Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (4). The standard errors that construct the 95-percent confidence intervals are clustered at the gj level.

Figure 5: Size of 1980 liberalization vs. the NTR gap



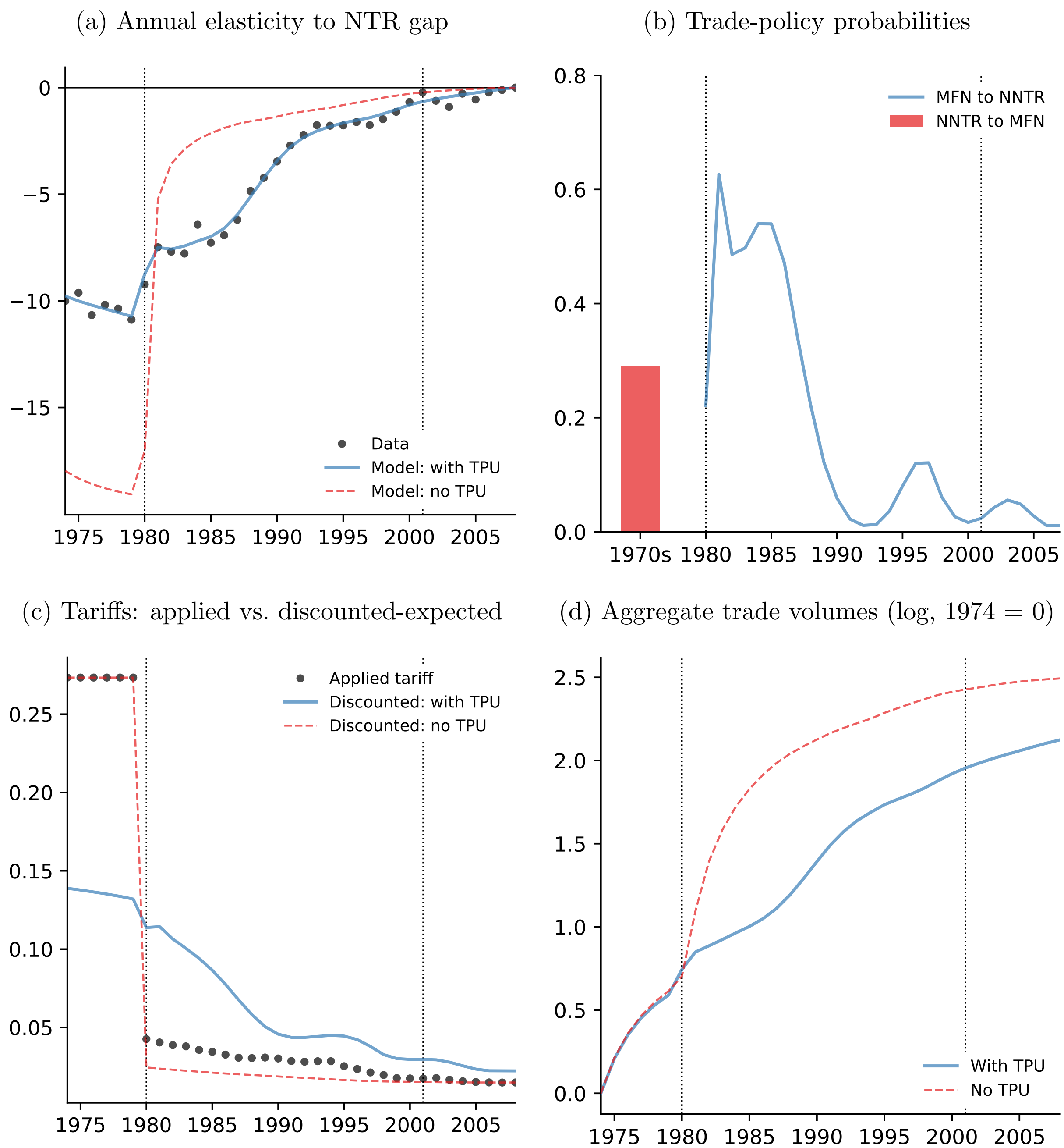
Notes: Each dot is a SITC good (excluding MFA goods). The slope of the fitted line (solid line) is the coefficient of regressing the change in applied duties on China's imports between 1979 and 1981 on the NTR gap (at the SITC level). The coefficient is 0.82 and its standard error 0.02. The R-square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

Figure 6: Sectoral Chinese exporter dynamics moments and the NTR gap



Notes: These figures plot the four moments used to discipline the exporter dynamics in the model against the average sectoral NTR gap. Exporter moments in panels (a)–(d) are also described in Table 3 and are obtained using firm-level data from Chinese manufacturers. This data is described in Appendix A. The demand elasticities are also reported in Column 1 of Table 4 and are obtained by averaging over the demand elasticities corresponding to U.S.-China from Soderbery (2018). The dashed lines are the regression coefficients of the mean NTR gap on the respective moment.

Figure 7: Trade and policy dynamics in the model with and without policy uncertainty

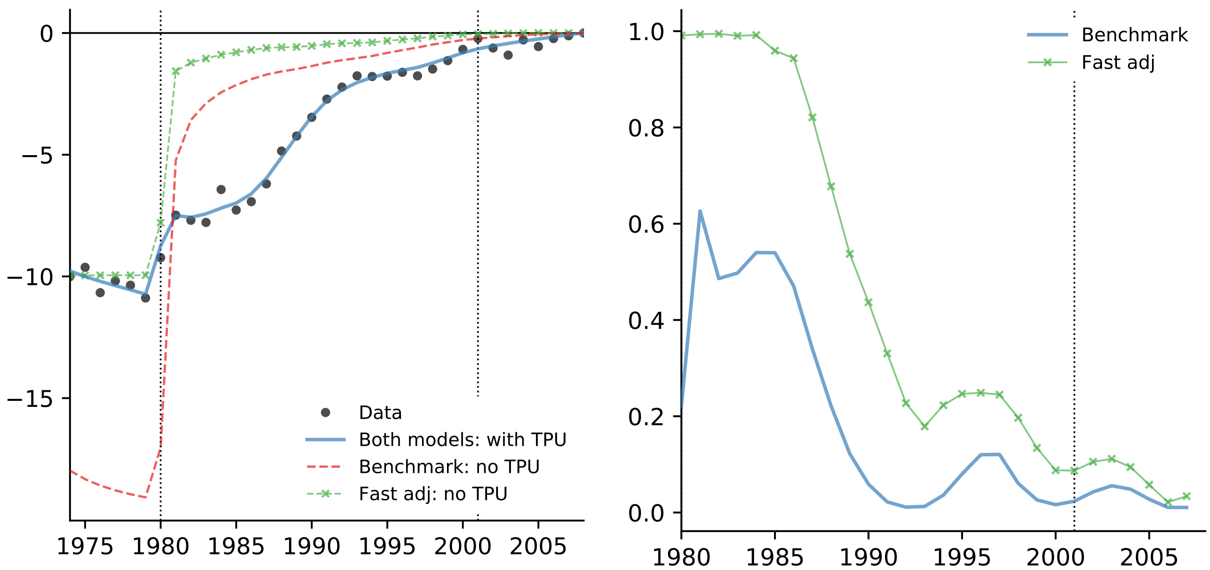


Notes: Panel (a) shows the NTR-gap elasticities estimated using simulated data from the model. Panel (b) shows the estimated probabilities of switching policy regimes. Panel (c) shows the discounted expected value of tariffs implied by these probabilities. Panel (d) shows the aggregate trade volumes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

Figure 8: Benchmark model vs. fast-adjustment model

(a) Annual elasticity to NTR gap

(b) Probability of losing MFN access

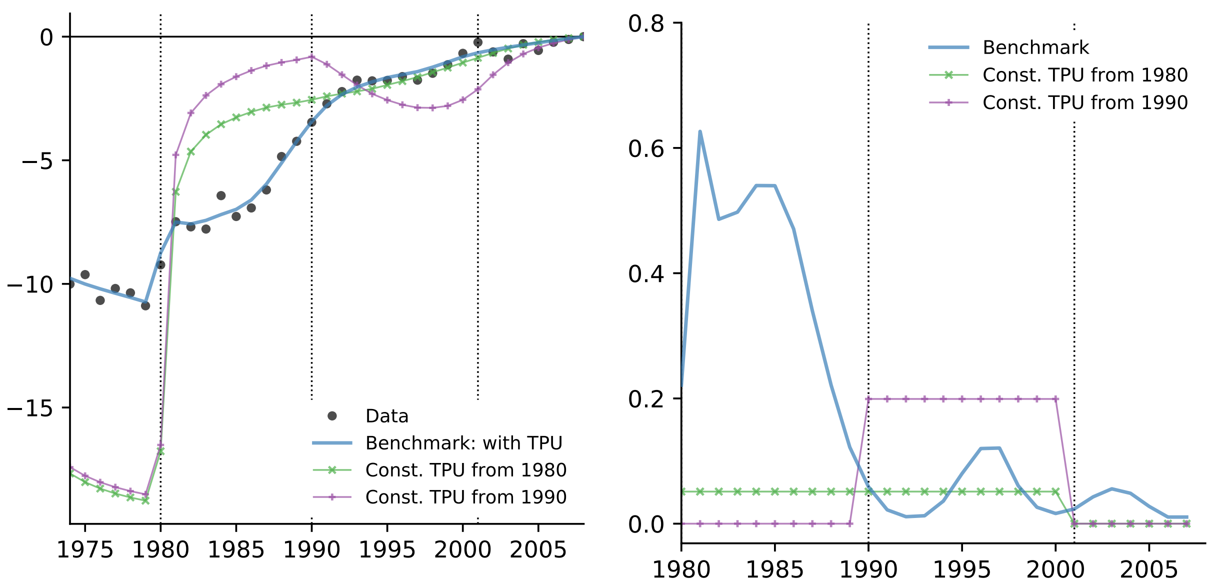


Notes: Panel (a) shows the NTR-gap elasticities estimated using simulated data from the model. Panel (b) shows the estimated probabilities of switching policy regimes. In the fast-adjustment model, there are no idiosyncratic shocks to firms' productivities or variable trade costs.

Figure 9: Benchmark model vs. constant-TPU models

(a) Annual elasticity to NTR gap

(b) Probability of losing MFN access

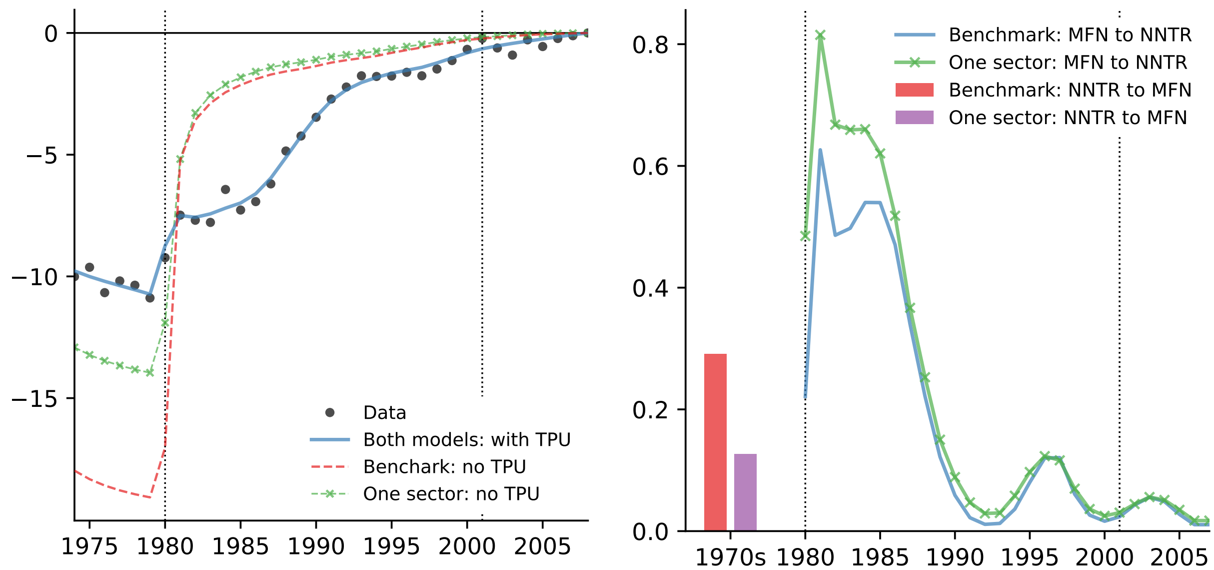


Notes: Panel (a) shows the NTR-gap elasticities estimated using simulated data from the model. Panel (b) shows the estimated probabilities of switching policy regimes. In the constant-TPU models, the probability of switching from the MFN regime to the NNTR regime is constant from either 1980 onward or 1990 onward.

Figure 10: Benchmark model vs. one-sector model

(a) Annual elasticity to NTR gap

(b) Trade-policy probabilities

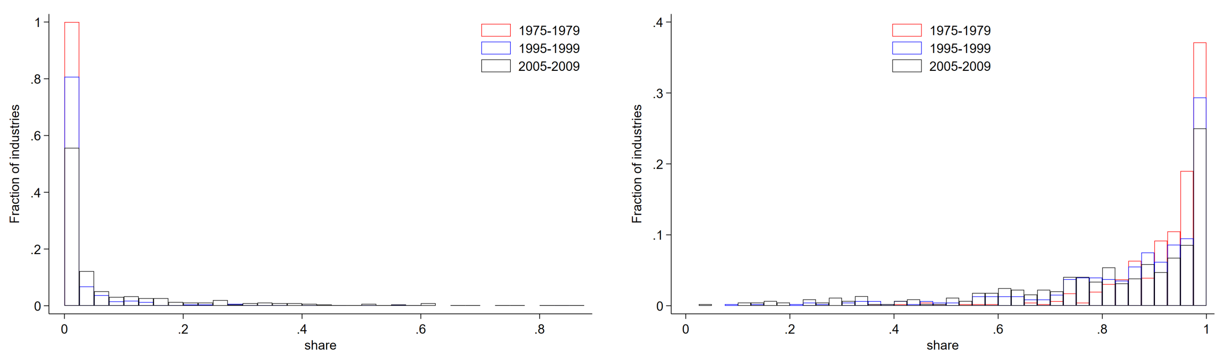


Notes: Panel (a) shows the NTR-gap elasticities estimated using simulated data from the model. Panel (b) shows the estimated probabilities of switching policy regimes. In the one-sector model, all goods have the same demand elasticity, non-tariff trade costs, and productivity dispersion, and these parameters are calibrated to match aggregate exporter-dynamics statistics instead of sector-level statistics.

Figure 11: Import export exposure across industries

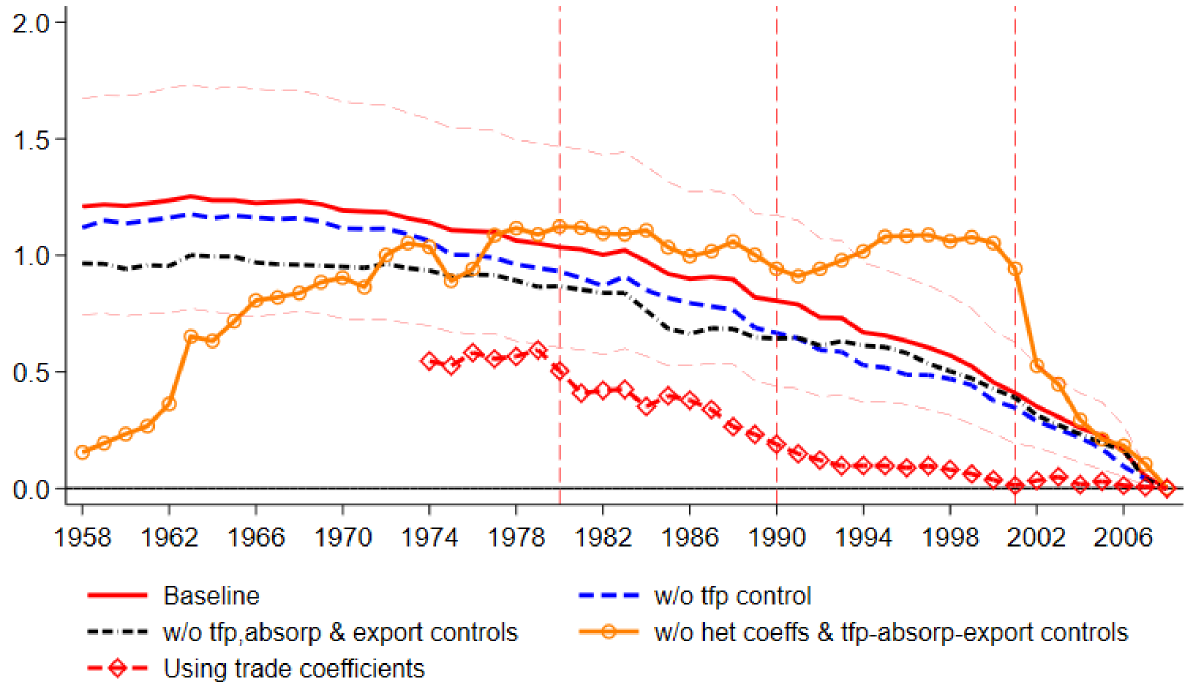
(a) Chinese import share

(b) Domestic sales share



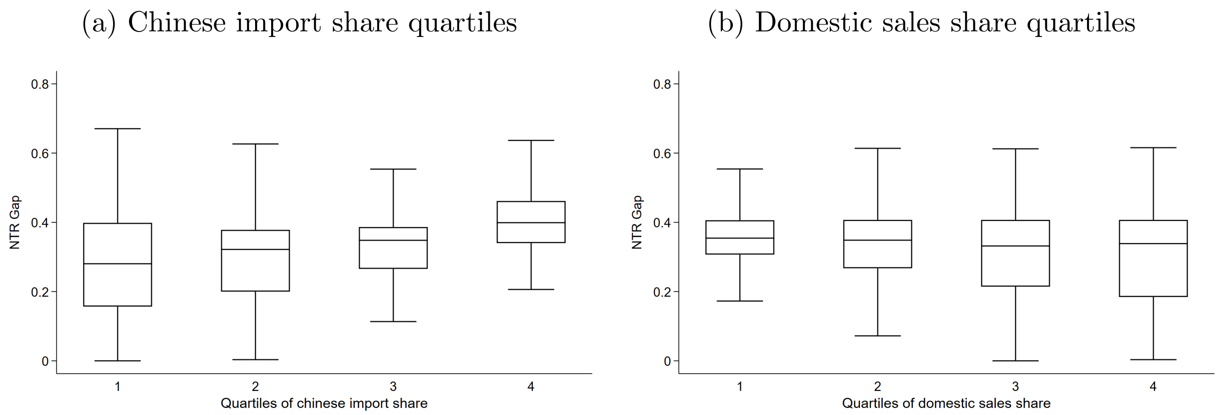
Notes: This figure plots the distributions of China's import share of total industrial absorption and the domestic market share of U.S. firms in total industry sales. The shares are simple averages over 5-year windows.

Figure 12: Sectoral employment effects of TPU reduction



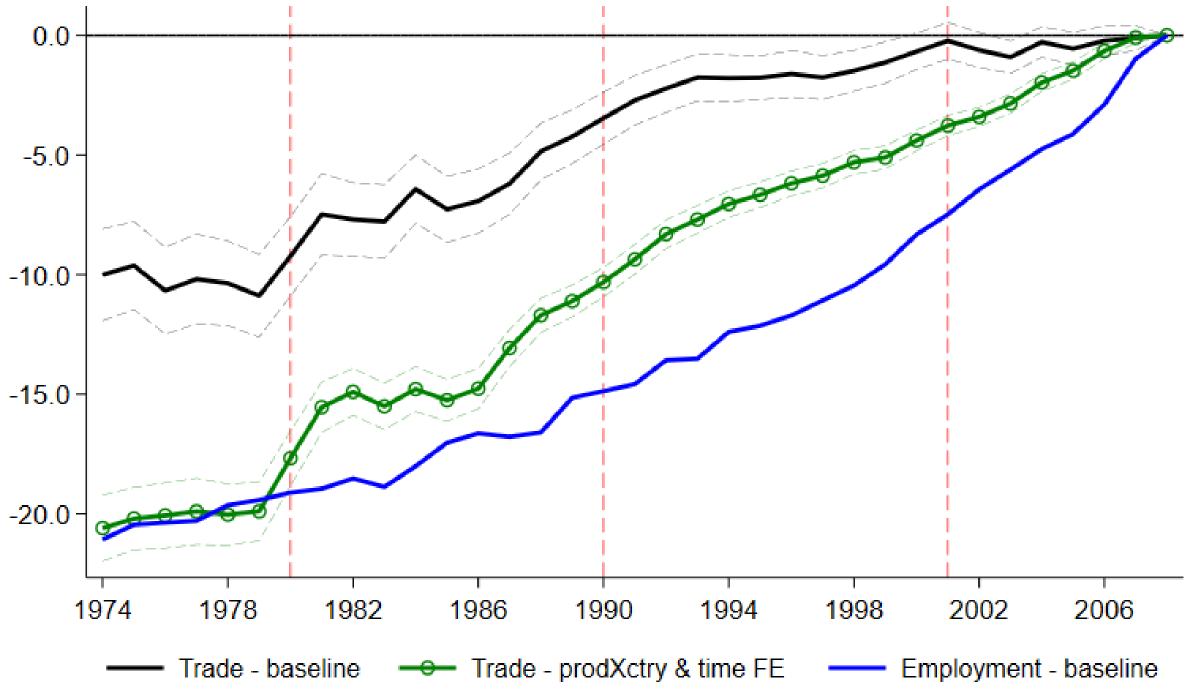
Notes: This figure plots the results from estimating (28). Since the annual coefficients are industry-specific, the baseline, in solid red, plots the industry with a median domestic sales share (91%) and the 90$^{\text{th}}$ percentile of the Chinese import share of domestic absorption during 1995–1999. This baseline equation is modified by removing control and heterogeneous coefficients to arrive at the orange line. Each regression follows Pierce and Schott (2016) and weights industries by their 1990 employment shares. The standard errors that construct the 95% confidence intervals around the baseline, in solid red, are clustered at the 4-digit SIC industry level.

Figure 13: Variation in the NTR gap



Notes: This figure plots the distribution of NTR gaps across different quartiles of China's import share of industry absorption (Panel a) and the domestic market share of U.S. firms in total industry sales (Panel b). Both are simple averages during 1995—1999.

Figure 14: Gap elasticity estimates from trade and employment data

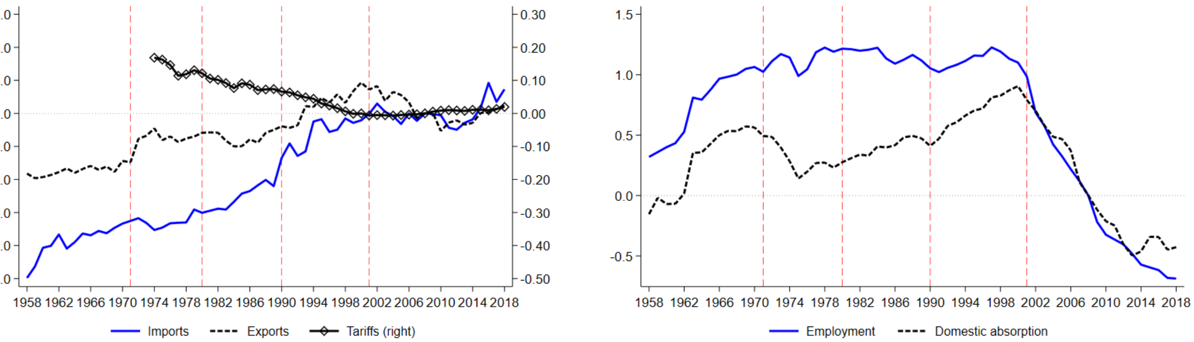


Notes: Employment - baseline plots the results from estimating (28) using the Chinese import share of domestic absorption and the industry export share during 1995–1999, using 1990 employment shares. *Trade - baseline* refers to our estimates from (4) with a full set of fixed effects for product-time (δ_{gt}), product-country (δ_{jg}), and country-time (δ_{jt}). *Trade - prodXctry & time FE* estimates the gap elasticity with a restricted set of fixed effects by product-country (δ_{jg}) and time (δ_t).

Figure 15: Elasticity to the NTR gap

(a) Total imports and exports

(b) Employment and domestic absorption



Notes: This figure plots the elasticity of aggregate variables to NTR gap over time. The tariffs in panel (a) are applied rates aggregated at SIC industry-year level by taking a weighted average across countries.

Appendix (For online publication)

In appendix A we discuss the firm level data from China used in our model calibration. In Appendix B, we document the robustness of our results concerning short- and long-run responses of trade to tariff changes and the gradual adjustment of U.S. imports from China. In Appendix C, we document the robustness of our results about the annual elasticity of trade to the NTR gap. In Appendix D, we present the results of additional quantitative experiments. In Appendix E we discuss some additional sensitivity relating U.S. sectoral employment to trade policy. Appendices F and G contain the additional tables and figures discussed in the aforementioned sections.

A Chinese firm-level data

The data for Chinese firms comes from an annual survey of manufacturing enterprises collected by the Chinese National Bureau of Statistics.[39] The dataset includes non-state firms with sales over 5 million RMB (about 600,000 U.S. dollars) and all state firms for 1998–2007. Information is derived from the balance sheet, profit and loss statements, and cash flow statements. The raw data consist of over 125,858 firms in 1998 and 306,298 firms in 2007 and includes sales, export revenues, value added, and number of employees. Firms are classified into industries according to the 4-digit Chinese National Industrial Classification (CNIC). To concord these with our goods classified under the SITC (revision 2) we proceed as follows. First, we apply the concordance between the 2-digit CNIC and the 3-digit ISIC (revision 2) reported in Table F.1, obtained from Xie et al. (2020). Next, we apply the concordance between the 3-digit ISIC (revision 2) and the 4-digit SITC revision 2.[40]

B 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

[39] This data has been widely used to study Chinese manufacturing growth between the late 1990s and 2000s (see, for example, Bai et al. (2023)). We thank Dan Lu for sharing the data with us.
[40] We obtain this concordance from Marc Muendler's website.

projection results are available upon request.

Shipping costs. Column 2 of Table F.2 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 F.2 report the ECM results under different specifications 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 F.3 presents the results of our ECM regression applied to more disaggregated datasets: 8-digit TSUSA and the HS-8. The former is available for 1974–1988, while the latter is available for 1989–2008. 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 classification, 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 TSUSA 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 gradual response of trade to changes in tariffs.

C Robustness and extensions: NTR-gap elasticity

The time-varying pattern of the effect of the NTR gap on China's exports to the United States shown in Figure 4 is robust to a range of alternative approaches. Before we report these results, we report the non-time varying average NTR-gap elasticities estimated under the approach of Pierce and Schott (2016) and decompose our baseline results into an extensive and intensive margin.

Pierce and Schott (2016) replication. In Section 2.4, we estimated the NTR-gap elasticities for each year between 1974 and 2007, relative to 2008. Instead, in their seminal article, Pierce and Schott (2016) consider the differences in trade pre- and post-China's WTO accession. Their estimating equation is:

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

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, (C.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 (C.1) when expanding the sample period back to 1974. This requires using 5-digit SITC goods instead of the HS-8 tariff lines used in Pierce and Schott (2016).

Table F.4 reports the results of several versions of this regression. In Column 1, we use the same 1992–2007 sample as in Pierce and Schott (2016) and estimate $\hat{\beta} = -0.9$.[41] In Column 2, we use the full sample period, 1974–2008, 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 NTR gap in 1999. Columns 3 and 4 use the statutory NNTR rate as measured by Feenstra et al. (2002) in place of the NTR gap, while Columns 5 and 6 use the average applied duty during 1974–1979 in our trade

[41] 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.

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 NTR-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.

Extensive margin effects. In our model, both the slow adjustment to past reforms and the trade dampening effect of uncertainty about future policy are almost entirely driven by firm entry decisions. While it is well-established in other contexts that the extensive margin plays an important role in the response of trade to policy changes, there are no data that would allow us to conclusively measure the role of this margin in the growth of Chinese exports to the United States over our full sample period (1974–2008). Nevertheless, we find suggestive evidence that firms' entry and exit decisions played a large role in the trade patterns we have documented.

We perform the following decomposition of imports into an intensive and extensive margin. For each country-good-year import flow, we compute a proxy for the extensive margin as the number of (product $\times$ U.S. district of entry) pairs, where products are defined at the most disaggregated level of observation (TSUSA during 1974–1988, HS-8 during 1989–2008). We define the intensive margin as the total import value divided by the number of (product $\times$ entry district) pairs. We then estimate the annual elasticity of each margin of trade to the NTR gap, as in equation (4) of the paper. Figure G.2 shows that the extensive margin clearly plays an important role in the evolution of the overall NTR gap.[42]

China supply effects. Our baseline approach controls for good-specific U.S. demand shocks

[42] We validate our extensive margin proxy using data from Colombia, in which we observe exports to the United States at the establishment-product level. Panel (a) of Figure G.3 shows that our extensive margin proxy is highly correlated with the number of firms per SITC good. Similarly, Panel (b) illustrates that our extensive margin proxy is closely related to, and, if anything, underestimates the count of firm-product pairs within each SITC good. It is important to note, though, that while these correlations indicate our approach is likely to yield a good proxy for the extensive margin, caution should be taken in interpreting the results in Figure G.2 as a quantitative decomposition of the contributions of the extensive and intensive margins to the overall NTR-gap elasticity.

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 the Chinese supply shocks were systematically correlated with the NTR 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 could have spillover effects and increase exports to other destinations, thereby introducing a downward bias in the elasticity estimates. We control for the importance of potentially confounding supply factors, using the World Trade Flows dataset from Feenstra et al. (2005) for 1974–2000 merged with the BACI trade database for 2001–2008 (Centre d'Études Prospectives et d'Informations Internationales, 2023).[43] We estimate

$$v_{ijgt} = \sum_{t'=1974}^{2007} \beta_t \mathbb{1}_{\{t=t' \wedge i=U.S. \wedge j=China\}} Gap_g + \delta_{igt} + \delta_{jgt} + u_{ijgt}, \tag{C.2}$$

where δ_{jgt} controls for exporter supply shocks and δ_{igt} controls for importer demand shocks (and trade barriers). Our coefficients of interest now include the difference in China's exports to the United States versus other destinations as well as the triple difference from our baseline approach. The results are shown in Figure G.4. The solid line plots the response of U.S. imports from China when considering U.S. imports only. This is analogous to (4), but uses the data in the world trade sample. The dashed line is the estimate when we include all trade flows, thus allowing us to include source-good-time fixed effects as, in (C.2). The overall pattern is similar to the baseline; in fact, the differences between the results from this specification and our baseline are not statistically significant.

Level of aggregation. In our baseline sample, we define goods at the 5-digit SITC (revision 2) product level. This allows us to construct a continuous dataset for 1974–2008 using well-established concordances between the underlying product schedules at which U.S. tariffs were determined, namely the TSUSA product schedule between 1974–1988 and the HS product

[43] 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. We 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.

schedule thereafter. To demonstrate that our results do not depend on product aggregation, we perform two robustness exercises. First, we estimate the NTR-gap elasticities using the product-level aggregation of the tariff lines of the two periods separately. Figure G.5 plots the results of (4) using the two tariff-line product classifications as well as the one obtained under the SITC classification. For 1989–2008, the results using the HS and the SITC classifications (Panel b) are nearly indistinguishable. In 1974–1988, the elasticities in the early years under the SITC classifications are slightly larger than under the TSUSA classification (Panel a). Qualitatively, however, they are very similar.[44] Second, we estimate (4) for the continuous sample period using a new concordance from Acosta and Cox (2022) between the TSUSA and the HS schedules.[45] The NTR-gap elasticities under this approach are reported in Figure G.6. Again, the path of the annual NTR-gap elasticities is similar.

Sample of countries and goods. Table F.7 demonstrates that our estimates of (4) are robust to our baseline sample design. The results are virtually unchanged when we include all countries (Column 2). When we relax the MFA exclusion (Column 3), the estimated annual elasticities of trade to the NTR gap fall by 10–20 percent, 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 4).[46] Finally, Column 5 shows the results when considering U.S. imports from China only. The results illustrate that not controlling for good-specific U.S. demand shocks leads to slightly larger elasticities in the early sample period.

Alternative NTR-gap measures. In our baseline specification (4) we consider the 1999 NTR gap as our measure of the 1980 liberalization/tariff risk to facilitate comparison with Pierce and Schott (2016). Recall that this gap is defined as the NNTR rate, established by the 1930 Smoot-Hawley Trade Act, minus the MFN rate in 1999, the year before China

[44]We also report robustness to these results in Table F.5 (TSUSA) and Table F.6 (HS-8).

[45]After accounting for the n-to-1, 1-to-n, and n-to-n relations we end up with around 2,800 unique product codes, 40 percent more than our baseline SITC classification.

[46]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 increase of around −1 percentage points in the elasticity of trade to the NTR gap.

joined the WTO. To illustrate that most of the variation is due to the NNTR rate we have estimated (4) defining Gap_g as the NNTR rate only. The results are reported in Column 1 of Table F.8. Using the NNTR rate yields annual gap elasticities about 10–20 percent smaller in the first years, but the gradual nature of the adjustment is unchanged. To show that the NTR gap is very related to the size of the 1980 liberalization (as already illustrated by Figure 5) we have also estimated (4) defining Gap_g as the average applied tariff rate to China during 1974–1979. The results yield elasticities that are about 5–15 percent smaller than in the baseline and converge to zero (or statistical insignificance) slightly faster in the last five years before PNTR access (Column 2 of Table F.8). Finally, we considered a time-varying version of the NTR gap, in which we define Gap_g as the NNTR rate minus the average tariff applied to all countries with NTR status that are not part of a free trade agreement with the United states. Again, the results are similar to our baseline and reported in Column 3 of Table F.8.

Additional trade-cost controls. Our baseline estimation (4) departs from Pierce and Schott (2016) by excluding applied tariff rates τ_{jgt} because the NTR gap is highly correlated with pre-1980 applied tariffs (Figure 5). Columns 4–5 of Table F.8 report the results when we include shipping cost and applied tariff rates. Both specifications leave the estimated coefficients virtually unchanged.

Anticipation 1979. Figure 1 shows that Chinese exports to the United States grew strongly in 1979, but Figure 4, which shows the elasticity to the NTR gap fell in 1979, indicates this increase was smaller for high-gap goods than low-gap goods. As emphasized by Khan and Khederlarian (2021), the weak growth for goods whose tariffs were about to decline the most could reflect anticipatory behavior by importers. To control for this possibility, we estimate a version of (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 6 of Table F.8 reports the results. Including the lead change smooths the response of trade flows to the NTR gap around 1979 (we no longer see a drop in the elasticity), but has no effect in other years.

D Robustness: Quantitative analysis

In this section, we conduct four 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 conduct a version of the no-TPU counterfactual in which the 1980 reform is anticipated, rather than unanticipated. In the last, we compare the results of our model to a model-free Bayesian learning exercise.

Bounds for our estimates of trade-policy probabilities. To estimate a path of expectations about future U.S. 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 NTR gap. In our empirical analysis, we also reported 95-percent 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, 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.[47] The results are shown in Figure G.7.

Our bounds for the probability of switching from NNTR to MFN tariffs (shown in light blue in the figure) are about 13 percentage points 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 14–17 percentage points 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

[47] 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.

China's WTO accession in 2001, the bounds for this probability are only a few percentage points 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.

Unanticipated changes in regime-switching probabilities. We have assumed that firms in our model know the entire path of probabilities of switching between trade policy regimes. In this section, 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 NTR gap coefficients while leaving all other aspects of the calibration unchanged. As in the first alternative, the realized path of tariffs is the same as in the benchmark, so any differences in outcomes are due to differences in firms' expectations.

Panel (b) of Figure G.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.

No-TPU counterfactual where 1980 reform is anticipated. We have assumed that in the no-TPU counterfactual, the 1980 reform is unanticipated. We think that this makes clearer the identification of the probability of gaining NTR status during the 1970s, but it is instructive to study what would have happened if the 1980 reform was anticipated.

Figure G.9 adds to our benchmark results a second no-TPU counterfactual in which, when the embargo is lifted in 1971, firms learn that NTR status will be granted in 1980 and that it will last forever. In this alternative counterfactual, the elasticity of trade to the NTR gap rises throughout the 1970s, accelerating as the 1980 reform nears. This is because many more firms producing high-gap goods begin exporting during the 1970s, as they know that tariffs on their goods will soon fall. The NTR-gap elasticity in this alternative counterfactual remains higher than in the benchmark counterfactual throughout the 1980s, but by 1990, the two counterfactuals are essentially the same. Importantly, this means that anticipation of the 1980 reform does not materially affect the extent to which gradual adjustment to this reform accounts for post-WTO growth in high-gap exports. The coefficient on the Pierce and Schott (2016) trade regression (20) is the same in both versions of the counterfactual.

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 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 at a rate roughly consistent with our estimates.

We focus on the probability of losing MFN status after the 1980 liberalization, $\omega(1, 2)$. We assume Bayesian agents have beta-distributed prior beliefs about this probability when the liberalization occurs in 1980:

$$p^{prior}(\omega(1, 2)|a, b) = \frac{\Gamma(a + b)}{\Gamma(a) + \Gamma(b)}\omega(1, 2)^{a-1}(1 - \omega(1, 2))^{b-1}. \tag{D.1}$$

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 $\omega(1, 2)$, 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 $b = 1, 2, \dots, 5$, 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 G.10 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 NTR-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.

E Additional results on employment effects

In this section, we report some additional results on the effects of trade policy and trade on employment. We start by showing our results are robust to using other measures of employment. We then show how employment depends on industry import and export shares.

We combine several datasets to conduct our analysis. We use the industry-year level trade data for 1958–1971 from Feenstra et al. (2002) and augment it with country-industry-year level data for 1972–2018 from Schott et al. (2008). For tariffs, we use Feenstra (1996) from 1972 to 1994 and add data for 1995–2018 using Schott et al. (2008). Figure 15(a) shows a minor break in the gap elasticity around the switch periods.

Figure G.11 plots the gap elasticity using several alternative measures and using our structural and reduced form equations. In the Panel (a), we show that the estimated effects on employment are a bit larger if we use production workers or production hours. The effects on sales are a bit smaller. Our model-consistent approach finds no effects prior to the lifting of the embargo in 1971. Panel (b) reports the gap elasticity using the reduced-form regression. All measures show a large increase from 1958 to 1974 and then a collapse around 2000. Most series peak in 1996.

Figure G.12 shows how the effects on employment depend on import penetration and domestic sales share. The Panel (a) shows that the effects on employment are concentrated in the sectors where China had the largest import penetration. The median sector is unaffected. Panel (b) shows that variation in the employment effects owing to domestic sales was modest in comparison to the variation from import penetration.

F Additional tables

Table F.1: Product-sector concordance

		CNIC	ISIC Rev.2
1	Food, beverage and tobacco	13-16	311-314
2	Textile, clothing, leather and footwear manufacturing	17-19	321-324
3	Wood and straw products	20	331-332
4	Paper and printing products	22-23	341-342
5	Energy products and chemicals	25-28	351-354
6	Rubber and plastic products	29-30	355-356
7	Non-metallic mineral products	31	361, 362, 369
8	Base metal manufacturing	32-33	371, 372
9	Calendered metal manufacturing	34	381
10	Other machinery and equipment manufacturing industry	35-36	382
11	Computer, electronic and optical products	40-41	385
12	Electrical equipment manufacturing	39	383
13	Vehicle manufacturing	37	384
14	Furniture and other manufacturing	21, 24, 42, 16	390, 332
15	Non-manufacturing	others	others

Notes: This concordance follows Xie et al. (2020).

Table F.2: Slow adjustment with ECM — robustness

	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***
	(0.78)	(0.79)	(0.68)	(0.77)	(0.67)	(0.82)
Long-run Others	−3.27***	−2.98***	−3.57***	−3.75***	−4.14***	−3.41***
	(0.22)	(0.22)	(0.20)	(0.20)	(0.18)	(0.26)
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
Adjusted R^2	0.27	0.28	0.27	0.26	0.26	0.26

Notes: The table reports estimates of (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 *Shipping* model includes shipping charges. The *All countries* 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 jg level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table F.3: Slow adjustment with ECM — robustness, continued

	1974–2008	1974–1988		1989–2008	
	SITC	TSUSA	SITC	HS-8	SITC
$\mathbb{1}\{j = China\}\Delta\tau_{jgt}$	-2.29***	-1.85***	-2.32***	-2.41***	-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***	-3.85***	-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***	-3.92***	-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***	-3.32***	-1.63***
	(0.11)	(0.15)	(0.21)	(0.16)	(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***	-7.58***	-6.00***
	(0.78)	(0.36)	(0.63)	(0.85)	(1.94)
Long-Run Others	-3.27***	-3.55***	-2.84***	-5.51***	-2.98***
	(0.22)	(0.19)	(0.31)	(0.26)	(0.28)
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	164,6275	613,403
Adjusted R^2	0.27	0.38	0.36	0.29	0.29

Note: This table reports estimates of (2) using the TSUSA and HS-8 product schedules for the available sample periods. For comparison, we also report the results using the SITC product classification for the respective sample periods. Column 1 reports our baseline estimate for the period 1974–2008 using 5-digit SITC products. Columns 2 and 3 use the sample period 1974–88, and Columns 4 and 5 1989–2008. Column 2 uses 7-digit TSUSA aggregation and Column 4 uses the 8-digit HS level. Standard errors in parentheses are clustered at the jg level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table F.4: NTR-gap elasticity — Pierce and Schott (2016) specification

	NTR Gap		Statutory NNTR		Applied NNTR	
$\mathbb{1}\{^{t>2000}_{j=China}\}NTRGap_g$	-0.92***	-2.50***				
	(0.22)	(0.28)				
$\mathbb{1}\{^{t>2000}_{j=China}\}NNTR_g$			-0.74***	-2.07***		
			(0.21)	(0.26)		
$\mathbb{1}\{^{t>2000}_{j=China}\}AppNNTR_g$					-0.84***	-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
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	839364
Adjusted R^2	0.84	0.78	0.84	0.78	0.86	0.80

Notes: Columns are estimates from variations of (C.1). Column 1 estimates the effect of the NTR gap on Chinese imports after PNTR access using the same period as Pierce and Schott (2016), but at the SITC aggregation level (see equation 5 in Pierce and Schott 2016). Column 2 uses our sample period from 1974 to 2008. Columns 3 and 4 estimates use the statutory NNTR rates instead of the NTR gap. Columns 5 and 6 use the applied NNTR rate calculated as the average applied rate on Chinese goods between 1974 and 1979 instead of the NTR gap. Standard errors in parentheses are clustered at the jg level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table F.5: NTR-gap elasticity — TSUSA, 1974–88

	Baseline	All Countries	Average NNTR, 1974-79	Tariffs
$\mathbb{1}\{_{j=China}^{t=t'}\}X_g$				
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	486,725	500,641	701,141	486,725
Adjusted R^2	0.80	0.80	0.80	0.80

Note: Columns are estimates of (4), except that (i) goods g are 7-digit TSUSA products, instead of the SITC products of our baseline; and (ii) the average 1974–1979 applied tariff on China is used as Gap_g instead of the NTR gap (not available for TSUSA). 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—*All Countries*— includes all countries. Column 3—*Average NNTR, 1974-79*— uses applied tariffs to all communist countries to calculate the NNTR rate, instead of applied tariffs to China only. Column 4—*Tariffs*—includes tariffs in (4) and, as expected, the coefficient diminishes in the early years. Standard errors are clustered at the jg level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table F.6: NTR-gap elasticity — HS-8, 1989–2007

	Baseline	No Tariffs	All Countries	Full	Balanced
$\mathbb{1}\{^{t=t'}_{j=China}\}X_g$					
1989	-3.42***	-3.45***	-3.36***	-2.25***	-3.36***
1990	-2.67***	-2.70***	-2.60***	-1.82***	-2.61***
1991	-2.18***	-2.22***	-2.14***	-1.55***	-2.15***
1992	-1.59***	-1.64***	-1.55***	-1.09***	-1.59***
1993	-1.03***	-1.08***	-0.98***	-0.83***	-1.07***
1994	-1.20***	-1.26***	-1.18***	-1.07***	-1.17***
1995	-1.06***	-1.10***	-0.99***	-1.00***	-1.25***
1996	-0.84***	-0.87***	-0.80***	-0.84***	-0.83***
1997	-1.01***	-1.03***	-0.99***	-0.98***	-1.04***
1998	-0.99***	-1.00***	-0.99***	-1.11***	-0.92***
1999	-0.63***	-0.63***	-0.62***	-0.89***	-0.62**
2000	-0.46***	-0.47***	-0.48***	-0.88***	-0.36*
2001	-0.50***	-0.52***	-0.51***	-0.99***	-0.34*
2002	-0.29*	-0.30*	-0.30*	-0.60***	-0.19
2003	-0.32**	-0.32**	-0.35**	-0.56***	-0.36**
2004	-0.33**	-0.32**	-0.32**	-0.54***	-0.22
2005	-0.31**	-0.31**	-0.31**	-0.07	-0.12
2006	-0.28**	-0.28**	-0.31**	-0.14	-0.23
2007	-0.17	-0.16	-0.19	-0.02	0.02
τ_{jgt}	-3.87***		-3.37***	-4.07***	-4.45***
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	2,023,284	2,023,284	2,314,918	2,930,184	1,214,290
Adjusted R^2	0.77	0.77	0.77	0.77	0.78

Note: All estimates are obtained using (4), except that (i) goods g are 8-digit HS products, instead of the SITC products of our baseline; and (ii) τ_{jgt} is included on 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—*No Tariffs*—excludes tariffs from the regression. Column 3—*All Countries*—includes all countries and Column 4—*Full*—further includes all goods. Column 5—*Balanced*—uses only products with non-zero U.S. imports from China in 1989 and/or 1990. Standard errors are clustered at the jg level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table F.7: NTR-gap elasticity — alternative samples

	Baseline	All Countries	Full	Balanced	Only U.S.-China
$\mathbb{1}\{_{j=China}^{t=t'}\}X_g$					
1974	-10.01***	-9.88***	-8.88***	-9.89***	-12.23***
1975	-9.62***	-9.54***	-8.37***	-9.50***	-12.38***
1976	-10.67***	-10.55***	-8.16***	-10.56***	-12.78***
1977	-10.19***	-10.10***	-8.00***	-10.08***	-11.82***
1978	-10.36***	-10.26***	-7.50***	-10.24***	-12.41***
1979	-10.89***	-10.79***	-7.64***	-10.78***	-12.51***
1980	-9.23***	-9.13***	-6.17***	-9.11***	-10.82***
1981	-7.48***	-7.38***	-4.83***	-7.50***	-8.91***
1982	-7.69***	-7.58***	-4.82***	-7.39***	-9.12***
1983	-7.78***	-7.67***	-4.48***	-7.30***	-9.21***
1984	-6.43***	-6.33***	-3.95***	-6.45***	-7.54***
1985	-7.27***	-7.18***	-4.63***	-6.77***	-8.19***
1986	-6.93***	-6.87***	-4.15***	-6.81***	-7.61***
1987	-6.20***	-6.09***	-3.82***	-6.11***	-6.73***
1988	-4.85***	-4.76***	-3.20***	-4.51***	-5.36***
1989	-4.23***	-4.08***	-2.54***	-3.80***	-5.08***
1990	-3.46***	-3.27***	-2.29***	-3.03***	-4.00***
1991	-2.72***	-2.60***	-1.73***	-2.53***	-3.08***
1992	-2.22***	-2.13***	-1.41***	-2.55***	-2.58***
1993	-1.76***	-1.63***	-0.90**	-2.07***	-1.90***
1994	-1.79***	-1.76***	-1.19***	-2.07***	-1.94***
1995	-1.77***	-1.67***	-1.18***	-1.85***	-1.87***
1996	-1.62***	-1.55***	-1.18***	-1.48***	-1.64***
1997	-1.76***	-1.78***	-1.17***	-1.78***	-1.93***
1998	-1.48***	-1.51***	-1.25***	-0.95**	-1.44***
1999	-1.13**	-1.19***	-0.88**	-1.23**	-1.19**
2000	-0.67*	-0.68*	-0.55	-0.86**	-0.49
2001	-0.23	-0.27	-0.21	-0.29	-0.06
2002	-0.62*	-0.63*	-0.29	-0.71*	-0.36
2003	-0.91***	-0.99***	-0.52*	-0.79**	-0.70*
2004	-0.28	-0.34	0.01	-0.31	-0.11
2005	-0.56	-0.64*	0.06	-0.26	-0.33
2006	-0.23	-0.37	0.31	-0.56*	-0.29
2007	-0.11	-0.14	-0.17	0.14	0.14
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	jt, gj
N	1,092,393	1,223,264	1,453,687	659,645	29,731
Adjusted R^2	0.78	0.78	0.78	0.80	0.78

Notes: This table reports the annual NTR-gap elasticities under alternative sample designs. All estimates are obtained using (4). Column 1 corresponds to our baseline estimates reported in Figure 4. Our baseline sample includes China and all countries with NTR, except when they enter a free trade agreement (e.g. Canada after 1989), and all goods except those included in the Multi-fiber Agreement (MFA). Column 2—*All countries*—includes all countries. Column 3—*Full*—further includes MFA-goods. Column 4—*Balanced*—includes only the goods with non-zero U.S. imports from China before 1981. Column 5—*Only U.S.-China*—includes only U.S. imports from China. Standard errors are clustered at the jg level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

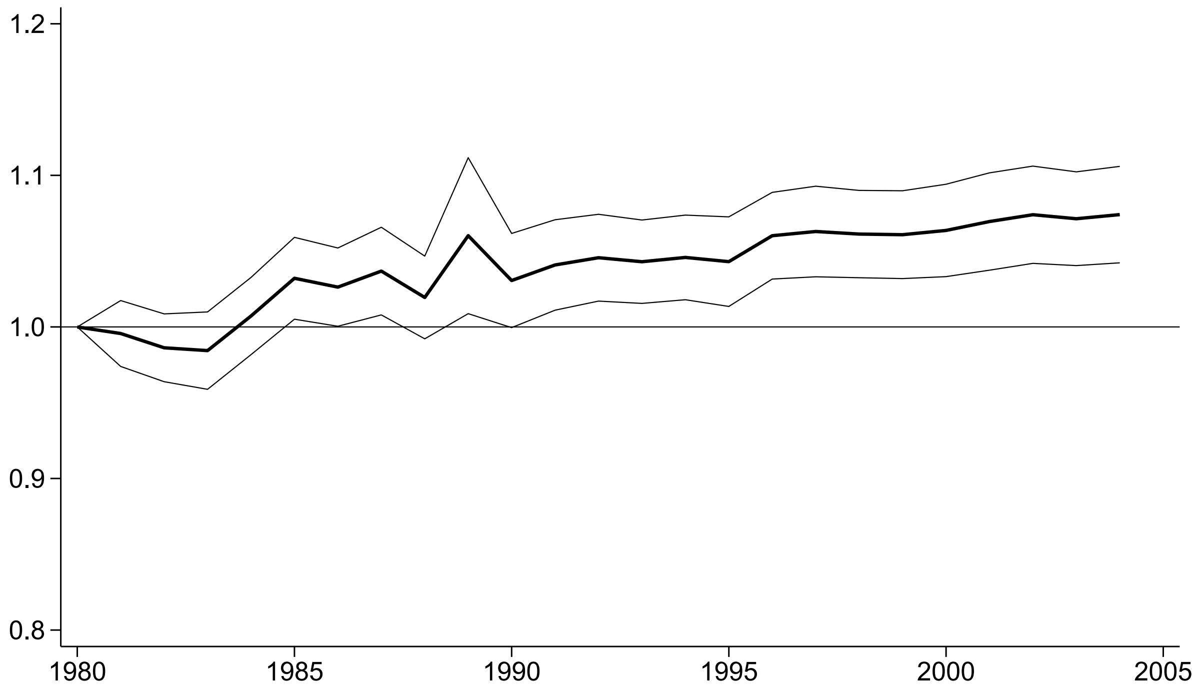
Table F.8: NTR-gap elasticity — additional robustness

	Alternative Gap measures			Other Trade Costs		
	NNTR	Applied	Time-Varying	Shipping	Tariffs	Anticipation
$\mathbb{1}\{^{t=t'}_{j=China}\}X_g$						
1974	-8.11***	-9.29***	-8.26***	-10.05***	-8.34***	-9.74***
1975	-8.31***	-9.31***	-8.66***	-9.66***	-7.95***	-10.09***
1976	-9.22***	-9.35***	-9.51***	-10.67***	-8.75***	-10.37***
1977	-8.99***	-8.88***	-9.31***	-10.22***	-8.34***	-10.42***
1978	-8.96***	-8.92***	-9.05***	-10.42***	-8.43***	-9.81***
1979	-9.44***	-9.18***	-9.71***	-10.92***	-9.06***	-9.83***
1980	-8.24***	-7.74***	-8.59***	-9.30***	-9.04***	-9.00***
1981	-6.28***	-6.64***	-6.52***	-7.51***	-7.34***	-7.39***
1982	-6.78***	-6.75***	-7.03***	-7.72***	-7.57***	-7.60***
1983	-6.64***	-6.33***	-6.88***	-7.77***	-7.68***	-7.39***
1984	-5.53***	-5.93***	-5.52***	-6.43***	-6.35***	-6.75***
1985	-6.55***	-5.90***	-6.54***	-7.26***	-7.18***	-7.07***
1986	-6.14***	-5.48***	-6.04***	-6.93***	-6.88***	-6.70***
1987	-5.59***	-5.09***	-5.31***	-6.25***	-6.21***	-5.99***
1988	-4.46***	-3.74***	-4.18***	-4.89***	-4.86***	-4.59***
1989	-3.69***	-2.79***	-3.31***	-4.13***	-4.10***	-3.97***
1990	-3.12***	-2.58***	-2.64***	-3.44***	-3.41***	-3.47***
1991	-2.29***	-2.01***	-1.75***	-2.79***	-2.74***	-2.47***
1992	-1.84***	-1.64***	-1.28**	-2.27***	-2.23***	-2.14***
1993	-1.37***	-1.58***	-0.78	-1.83***	-1.77***	-1.64***
1994	-1.39***	-1.52***	-0.83	-1.73***	-1.67***	-1.61***
1995	-1.50***	-1.09**	-0.94*	-1.77***	-1.74***	-1.71***
1996	-1.36***	-1.06**	-0.78	-1.62***	-1.54***	-1.52***
1997	-1.47***	-1.15**	-0.89*	-1.67***	-1.66***	-1.55***
1998	-1.39***	-0.88**	-0.78	-1.40***	-1.39***	-1.49***
1999	-0.92**	-1.07**	-0.28	-1.09**	-1.08**	-1.09***
2000	-0.56	-0.67	0.09	-0.70*	-0.70*	-0.32
2001	-0.23	-0.23	0.47	-0.30	-0.30	-0.30
2002	-0.52	-0.33	0.17	-0.62*	-0.62*	-0.57
2003	-0.83**	-0.55	-0.12	-0.85**	-0.85**	-0.65**
2004	-0.21	0.13	0.49	-0.27	-0.26	-0.37
2005	-0.63*	0.32	0.04	-0.55	-0.55	-0.48
2006	-0.19	-0.33	0.51	-0.22	-0.21	-0.30
2007	-0.15	0.29	0.58	-0.17	-0.17	0.00
log Shipping Cost $_{jgt}$				-3.03***	-3.02***	
τ_{jgt}					-2.40***	
$\Delta\tau_{jg,t+1}$						1.01***
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	1,092,393	617,167	1,092,393	1,092,393	1,092,393	870341
Adjusted R^2	0.78	0.80	0.78	0.79	0.79	0.79

Notes: This table reports the annual NTR-gap elasticities estimated by (4) under alternative specifications of the gap measure, additional trade cost controls, and control for anticipatory effects. Column 1—*NNTR*—defines Gap_g as the NNTR rate. Column 2—*Applied*—defines Gap_g as the applied NNTR rate, calculated as the good-level tariff rate applied to Chinese imports between 1974 and 1979. Column 3—*Time-Varying*—defines Gap_g as the NNTR rate minus the average duty applied to NTR countries in every year. Column 4—*Shipping*—includes shipping costs. Column 5—*Tariffs*—further includes applied duties. Column 6—*Anticipation*—includes the lead change in tariffs to control for some of the anticipation to the 1980 NTR liberalization. Standard errors in parentheses are clustered at the jg level. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

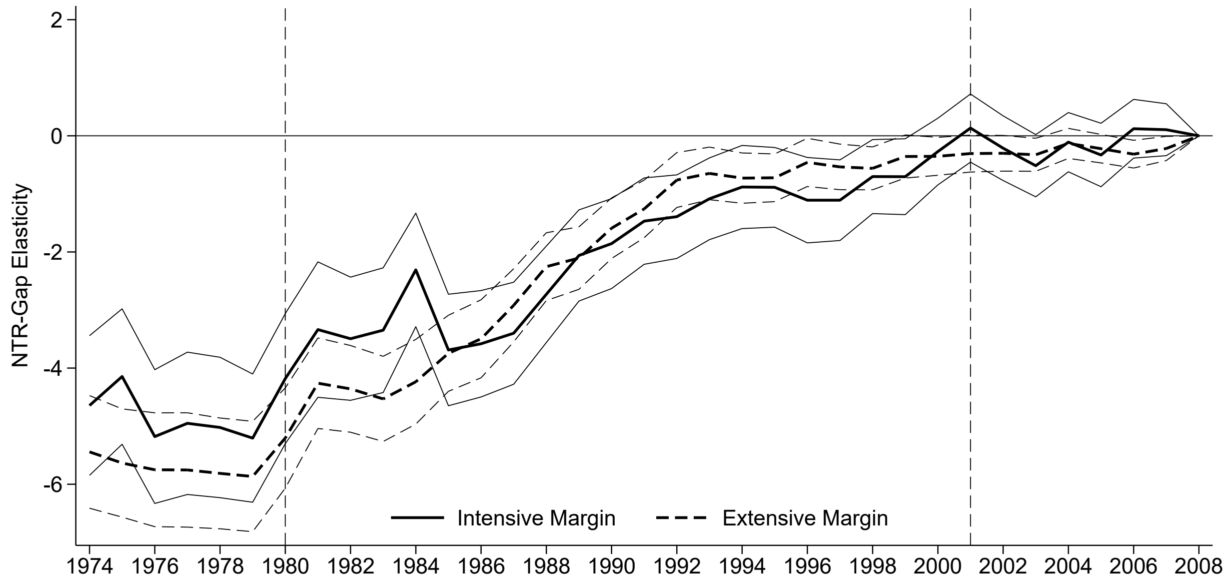
G Additional figures

Figure G.1: Autocorrelation of 1980 tariff changes



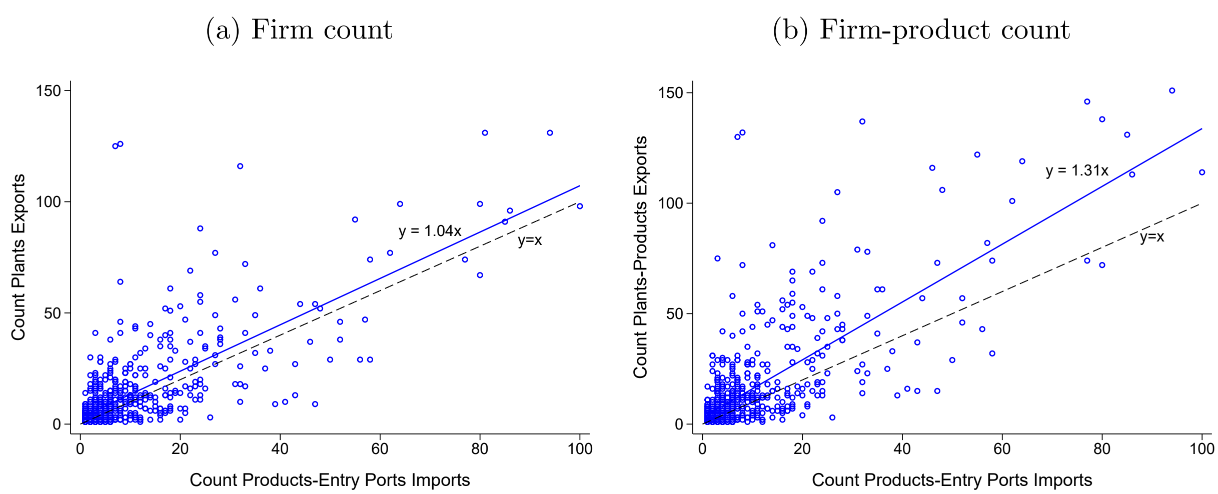
Notes: This figure plots the autocorrelation of the 1980 tariff changes, obtained from regressing the h-year ($h = [1, 25]$) tariff change relative to 1979 on the one-year tariff change between 1980 and 1979. The estimation includes the same fixed effects as in (3). The 95-percent confidence interval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

Figure G.2: Extensive and intensive margin trade elasticities to NTR gap



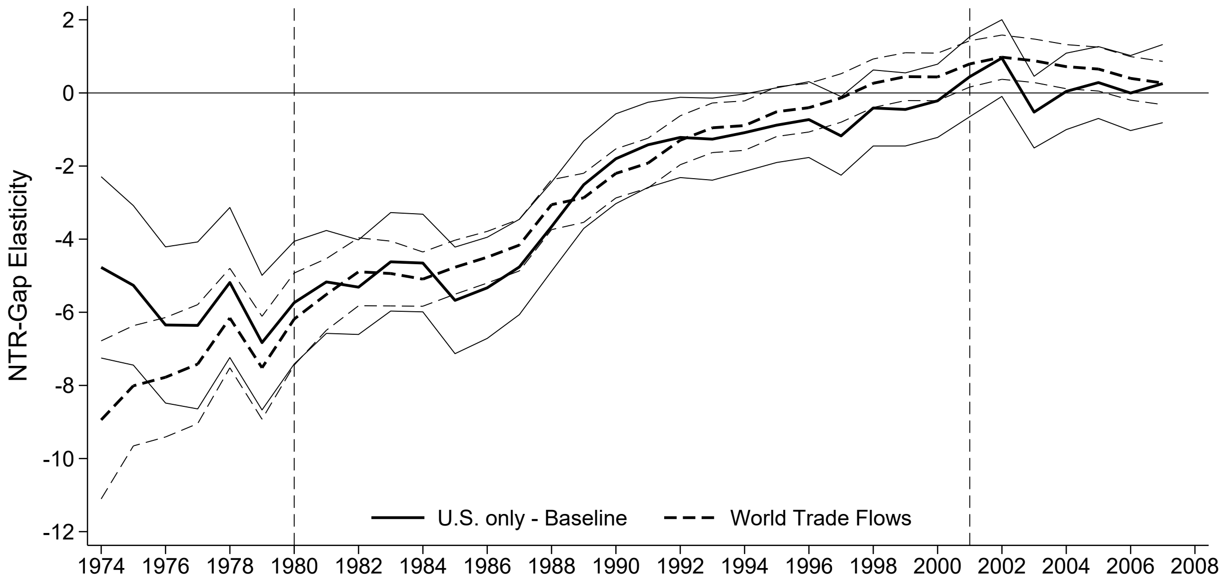
Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (4) using measures of the extensive (solid line) and intensive margins (dashed line) as the dependent variable. The standard errors that construct the 95-percent confidence intervals are clustered at the gj level.

Figure G.3: Extensive margin measures in Colombian firm-level export data



Notes: In both figures the x-axis is the extensive margin measure we use in Figure G.2 applied to U.S. imports from Colombia in 2006: the count of product-district of entry of U.S. imports, where products are defined at the TSUSA 7-digit (1974–1988) and HS 8-digit (1989–2008) level, of each SITC good. In Panel (a) the y-axis is the count of firms exporting each SITC good to the United States in the Colombian export data and in Panel (b) it is the count of firms-products of each SITC good.

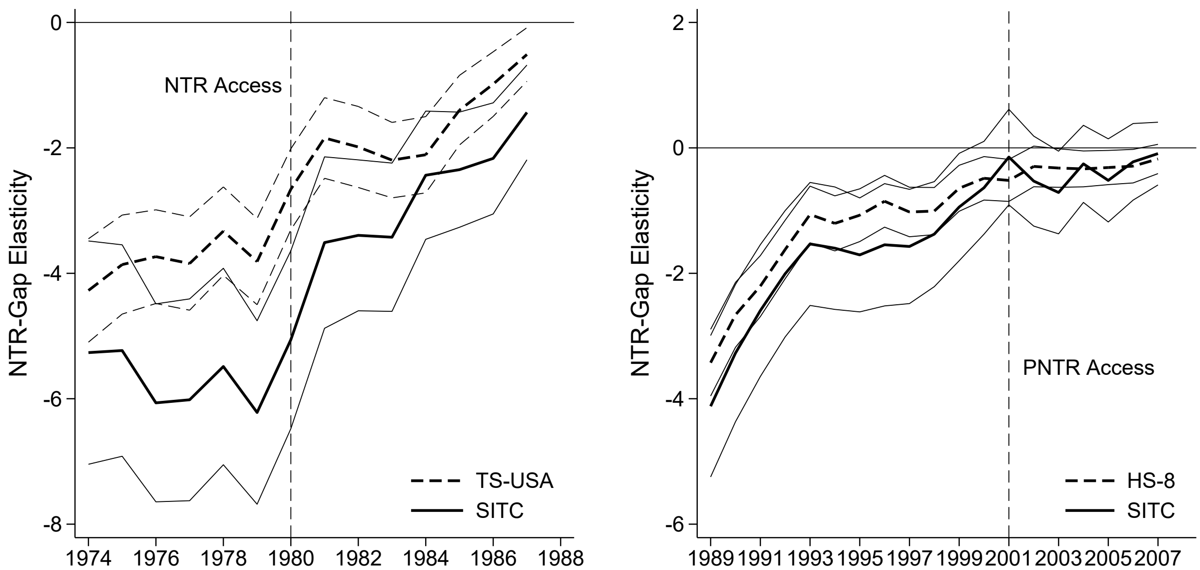
Figure G.4: Annual NTR-gap elasticities: China supply factors



Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (C.2) using the merged World Trade dataset from Feenstra et al. (2005) (1974–2000) and the BACI Trade Dataset (2000–2008). The standard errors that construct the 95-percent confidence interval are clustered at the jg level.

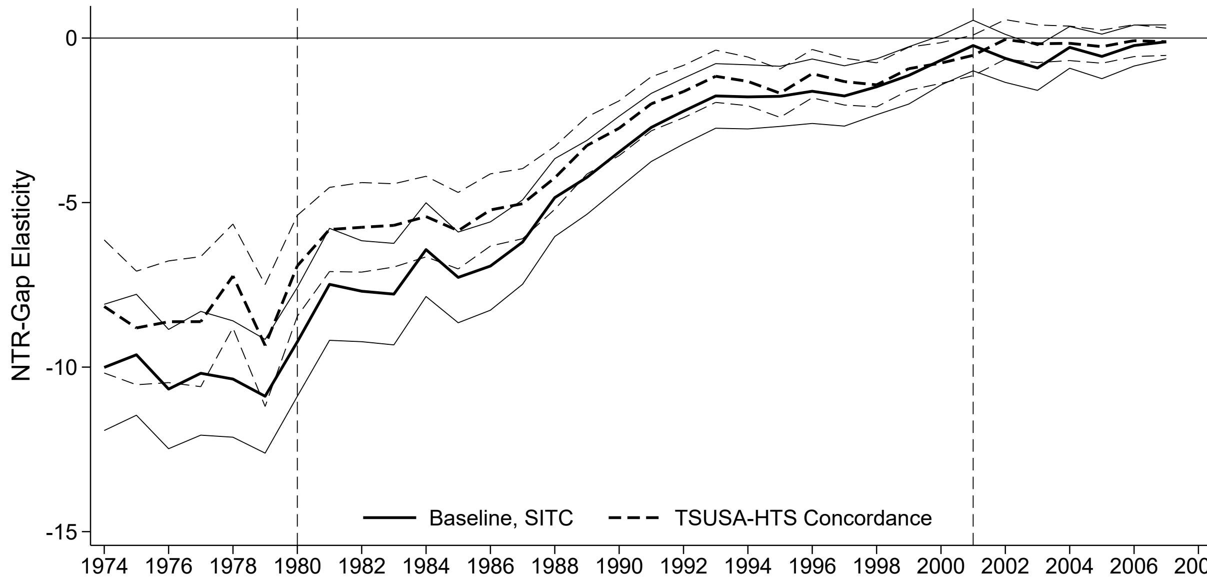
Figure G.5: Annual gap elasticities with aggregation at tariff lines

(a) TSUS 1974–1988 (b) HS 1989–2008



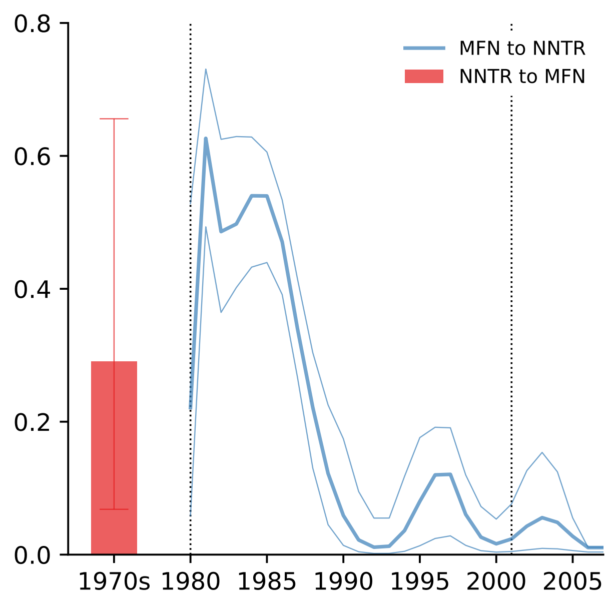
Notes: Both figures report estimates of (4) using the two product classification at which U.S. tariffs are set. This corresponds to the 7-digit TSUSA classification between 1974 and 1988 and the HS classification between 1989 and 2008. In both cases, we set the reference period to be the last year of the sample period (given our fully saturated specification of the fixed effects). For comparison, we also estimate (4) with our baseline product classification, the 5-digit SITC goods, for the corresponding sample periods. The standard errors that construct the 95-percent confidence interval are clustered at the jg level.

Figure G.6: Annual NTR-gap elasticities: SITC and TSUSA-HS



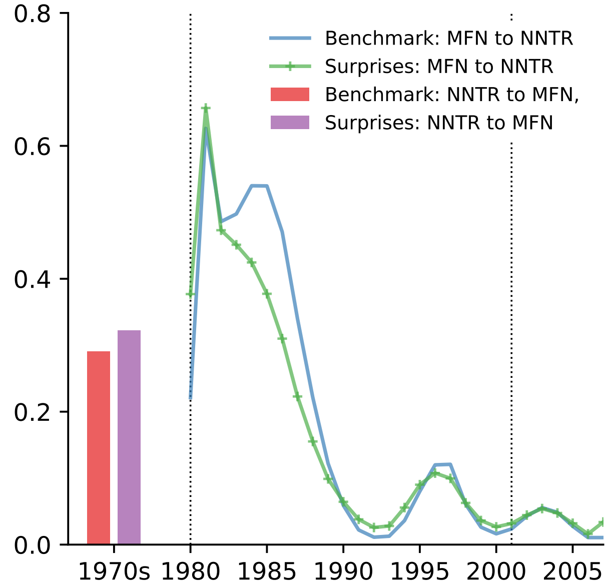
Notes: This figure plots our baseline (solid line) estimate of the annual NTR-gap elasticity obtained by (4) and the one (dashed line) using the concordance between TSUSA and the HS product schedules from Acosta and Cox (2022). The standard errors that construct the 95 percent confidence interval are clustered at the jg level.

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



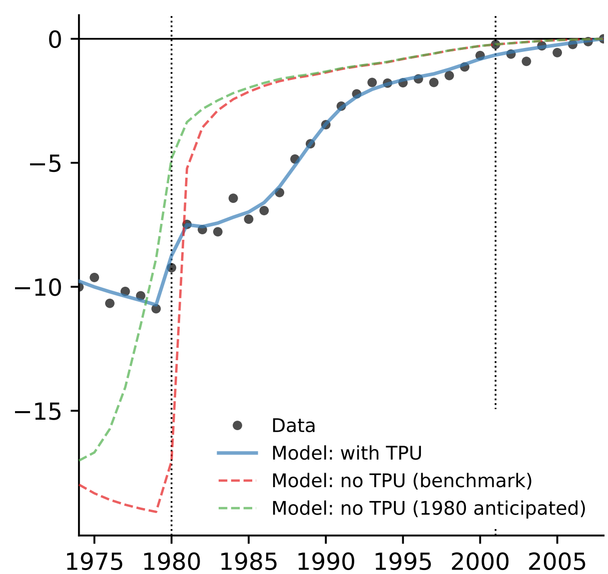
Notes: This figure shows the estimated probabilities of switching policy regimes. Thick lines are the baseline results. Thin lines are estimated by matching the upper and lower confidence intervals shown in Figure 4.

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



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

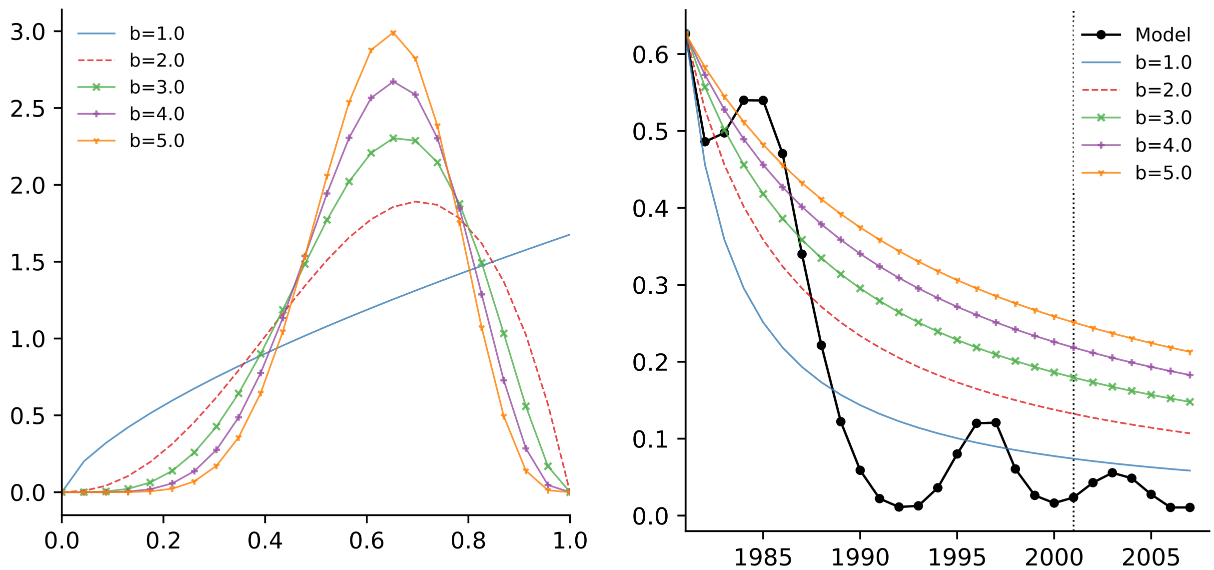
Figure G.9: No-TPU counterfactual: benchmark vs. anticipated 1980 reform



Notes: This figure plots the NTR-gap elasticities estimated using simulated data from the model. No TPU (benchmark): 1980 reform is unanticipated. No TPU (1980 reform anticipated): firms learn about 1980 reform immediately in 1971.

Figure G.10: Model-implied expectations vs. Bayesian learning

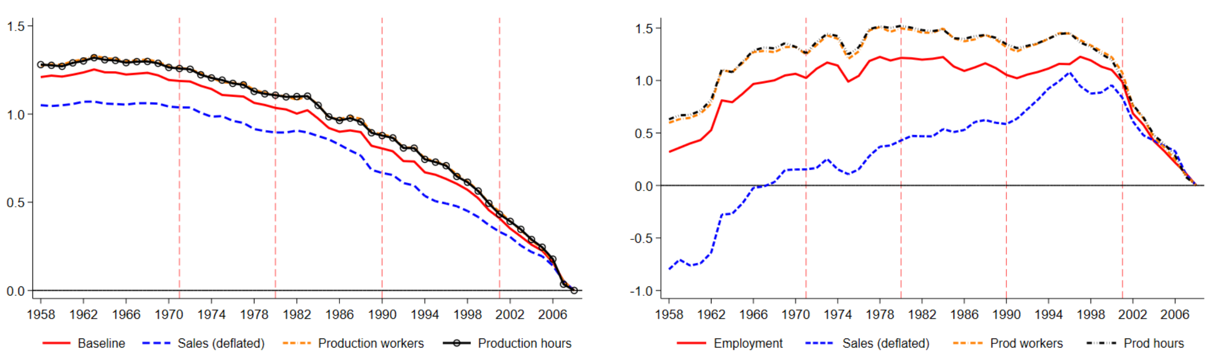
(a) Prior PDF (b) Posterior mean



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.

Figure G.11: Alternative measures of substitution

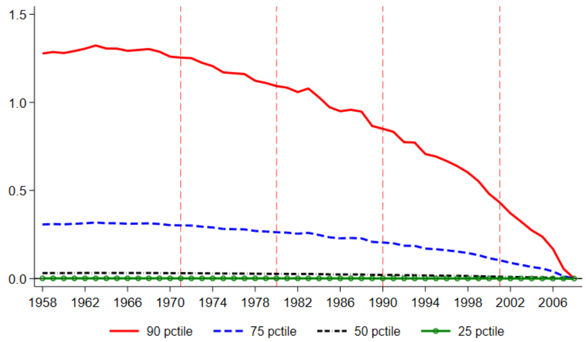
(a) Baseline estimate (b) Reduced form elasticities



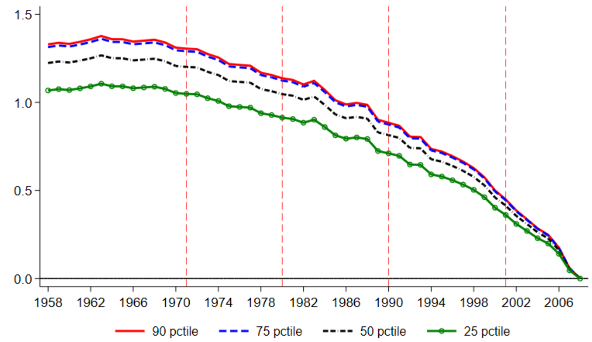
Notes: This figure plots versions of Figure 12 with sectoral deflated sales, production workers, and sectoral production hours as the dependent variables instead of total sectoral employment.

Figure G.12: Sectoral employment effects by industry import and export share

(a) Chinese import share percentiles



(b) Domestic sales share percentiles



Notes: This figure highlights the industry-level heterogeneity in the annual coefficients from (28). Panel (a) holds the domestic sales share constant at the median during 1995–1999 (91 percent) and varies the Chinese import share from different points in the distribution. Similarly, Panel (b) holds the Chinese import share constant at the median during 1995-1999 (6 percent) and varies the domestic sales share from different points in the distribution.