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

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Abstract

We study China's export growth to the United States from 1950–2008, using a structural model to disentangle the effects of past tariff changes from the effects of changes in expectations of future tariffs. We find the effects of China's 1980 Normal Trade Relations (NTR) grant lasted past its 2001 accession to the World Trade Organization (WTO), and the likelihood of losing NTR status decreased significantly during 1986–1992 but changed little thereafter. U.S. manufacturing employment trends support our findings: industries more exposed to the 1980 reform have shed workers steadily since then without acceleration around China's WTO accession.

JEL Classifications: F12, F13, F14

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1 Introduction

International trade depends on past, present, and future trade policy, but the three are rarely studied together. Many papers study 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 trade responds gradually to past policy changes, whereas another shows uncertainty about future policy can affect trade in the present. We argue the effects of past and future policy are tightly intertwined. We then 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 on China evolved and to measure how these expectations, together with changes in tariffs, shaped Chinese export growth. Different from previous studies that focus on the 1990s and 2000s, we take a longer-term view that begins when the United States lifted its embargo on China in 1971. We find in the 1970s, when the United States levied high Non-Normal Trade Relations (NNTR) tariffs on China, the probability that China would gain Normal Trade Relations (NTR), and face lower NTR tariffs, was about 30 percent. This uncertainty increased exports of goods with large gaps between the two tariff schedules, which we refer to as high-gap goods. After China gained NTR status in 1980, the probability that this status would be revoked was high, which depressed exports of these same high-gap goods. This reform became more credible in the late 1980s and early 1990s, and the change in expectations during this period led to sustained growth in Chinese exports to the United States.

Our key findings are that trade-policy uncertainty (TPU) 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 Pierce and Schott (2016) and Handley and Limão (2017), who interpret the faster growth of exports of high-gap goods relative to exports of low-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 fast: high-gap exports grew ten times faster than low-gap exports during 1985–1993 than during 2000–2006. Second, we find much of the 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, both to past reforms and to changes in expectations.

Two aspects of U.S. trade policy on China make it ideal for studying the dynamics of TPU. First, the United States has 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 1950–1971. 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 measures as 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) and a local-projections specification (Jordà, 2005), we find the long-run response of Chinese exports to U.S. tariffs is almost six times the short-run response and that completing 90 percent of the adjustment takes as long as 19 years. The second measurement is an estimate of how trade responds to 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. 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 1974–2007. These results indicate the effects of policy uncertainty were the largest in the 1970s and early 1980s and had largely dissipated before China joined the WTO.[1]

On their own, these reduced-form estimates do not identify causal effects or structural parameters. 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 goods experienced the largest tariff reductions when NTR status was first granted, and thus, they took the longest to converge to their new levels of trade. We overcome these identification challenges using indirect inference to estimate the time-varying transition probabilities between the 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 Das et al. (2007) that captures the key features of marginal exporter dynamics. Firms in each industry differ in productivity and variable export costs, which they reduce gradually through a risky investment.[2] 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 to match our estimates of the adjustment 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

[1] That the effect of the NTR gap on trade was largest in the 1970s before China had NTR status, suggests the NTR gap captures something besides exposure to tariff risk.

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

model begins in 1971 in the NNTR regime and switches to the NTR regime in 1980. We choose the probabilities of switching between trade-policy regimes 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 it was initially high, peaking at 59 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 agreement that would become the WTO, the probability of losing NTR status fell rapidly. It temporarily rose again in the early to mid 1990s, but by the late 1990s, it had fallen to five percent. Joining the WTO had a minor effect on the probability of losing access to NTR tariff rates; this probability fell by less than two percentage points between 1999 and 2007.

We isolate the roles of TPU and gradual adjustment to tariff changes in a model without 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 faster than in the benchmark model (and in the data). However, the gradual adjustment to the liberalizations in the counterfactual model is ongoing even after China's 2001 WTO accession. We find that gradual adjustment accounts for almost 45 percent 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 and past policy changes. To illustrate this point, we construct a counterfactual model in which the probability of losing NTR status is 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. Thus, 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 high-gap goods exports 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.

Finally, we revisit the effects of TPU on U.S. labor markets. 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 by 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 U.S. employment has declined as a result of changes in expectations about U.S. trade policy toward China. By contrast, 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 three 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. 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 tariff elasticities. We contribute by studying the dynamic response to a single exogenous, immediate tariff reduction over a multi-decade span using disaggregated data. Different from these studies, we do not interpret our results causally or structurally. Instead, we use these results as inputs to our indirect inference exercise, which uses a model designed to capture the main potential sources of bias in the data.

The second strand studies how expectations about 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- and industry-level data to identify the effects of TPU. This literature largely focuses on U.S. trade policy toward China.[3] Pierce and Schott (2016), Feng et al. (2017), Handley and Limão (2017), and Bianconi et al. (2021) measure the growth in trade, employment, and other outcomes that resulted from the elimination of uncertainty when China was granted PNTR status. Our contribution is to study how these tariff gaps, which also capture the size of the initial 1980 liberalization, influenced outcomes from the beginning of the U.S.-China trade relationship. Again, we use these results as inputs to our model, rather than interpreting them causally.

The third strand uses models to estimate trade-policy expectations. 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 to estimate the probability of NNTR reversal. We estimate a time-varying probability over a longer interval using a richer model in which tariff changes have persistent effects. Our analysis highlights the importance of earlier changes in trade-policy expectations during the late 1980s and

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

early 1990s in explaining the growth of Chinese exports to the United States around the turn of the century. We find ignoring these changes leads to an overstatement of the effect of PNTR access on tariff risk. Alessandria et al. (2017) estimate the expected path of inward and outward trade policy for China from macroeconomic time series. Complementary to our approach, Alessandria et al. (2024c) estimate a time-varying probability of NNTR reversal from 1990–2005 using within-year variation in trade flows and trade-policy risk in an s-S inventory model, obtaining similar results for this later period. More generally, our work is related to the classic "Peso problem" of estimating the likelihood of discrete events that are unobserved for long periods of time.

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, we estimate short- and long-run elasticities of trade to the 1980 tariff reduction and the speed of adjustment between the two horizons. Second, we study the elasticity of trade to the risk of reversing the tariff reduction and its evolution. Neither of these approaches directly identifies structural parameters or causal effects, because they cannot disentangle the effects of slow adjustment to past policy changes from uncertainty about future policy. However, they provide reduced-form evidence that these margins are important. They also produce crucial inputs to our quantitative analysis.

2.1 Data

We use annual data on U.S. imports from 1974–2008, 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.[4] We refer to this level of aggregation as a *good* and denote it by g. Our sample contains 1,469 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. Applied tariffs, denoted τ_{jgt}, are applied duties divided by FOB import values. Ad-valorem-equivalent NNTR and NTR statutory tariffs, from Feenstra et al. (2002), at the 8-digit level of the Harmonized Tariff Schedule (HS), are mapped to the 5-digit SITC classification. The SITC-level NNTR and NTR tariffs (τ_g^{NNTR} and τ_{gt}^{NTR}, 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 NTR 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 was not part of preferential trade agreements with the United States, excluding Canada, Mexico, and other communist countries.[5] A key feature of our sample is that all the countries, including China, faced approximately the same tariffs after 1980.[6] 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

[4]More disaggregated 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.

[5] Countries excluded that held NNTR status 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 Soviet Union.

[6]Variation in bilateral applied tariffs among NTR countries is due to aggregation, specific tariffs, or measurement error. This variation, however, is minor.

accession to the WTO triggered the removal of import quotas on these goods, and our tariff 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 1950 and 2018, three key reforms to U.S. tariffs on Chinese goods were enacted: lifting the embargo in 1971, granting conditional NTR status in 1980, and granting permanent NTR status 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 faced the relatively high NNTR tariff rates set by the Smoot-Hawley Act of 1930. From 1980 until the 2018 trade war, Chinese goods faced the much lower NTR tariffs applied to imports from WTO members (and non-members that, like China, have been unilaterally granted NTR status). In the appendix, we summarize the NNTR and NTR tariff schedules at the 2-digit level of the Chinese Industrial Classification System.[7] The mean NNTR rate is 29 percent with a standard deviation of 21 percent. The average applied tariff for non-China NTR countries was six in 1979 and two in 2001. For China, the average applied rate was considerably higher in 1979 (28 vs. six), but identical by 2001 (two vs. two). Figure 1 shows the evolution of the distribution of applied tariffs on Chinese goods. The median tariff fell from 35 percent to about 8 percent. The vast majority of this decline occurred in 1980. Subsequent tariff reductions in NTR tariffs were related to gradual phaseouts from successive rounds of the General Agreement on Tariffs and Trade (GATT).

The end of the embargo and the 1980 liberalization were abrupt events, unlike the GATT rounds, which featured gradual tariff phaseouts that were agreed upon in advance. Figure 2 plots the distribution of the residuals from regressing annual tariff changes on country-year, good-year, and country-good fixed effects. Virtually all the variation in tariff changes over

[7]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.

and above multilateral changes in NTR tariffs, which are absorbed by the fixed effects, was due to the granting of NTR status in 1980.[8]

U.S. trade policy toward China was uncertain, and the extent of that uncertainty varied with domestic and international politics. After the embargo was lifted in 1971, what path would lead to the United States granting China NTR rates was unclear. A series of trade acts in 1951, 1962, and 1974 required imports from non-market economies to face NNTR rates. Exceptions to this requirement were rare, typically contingent on substantial reforms, and could be removed quickly (or withdrawn before being implemented), eliminating this requirement was discussed periodically.[9] Further uncertainty arose from the lack of diplomatic relations between China and the United States.[10]

When China was granted temporary NTR status in 1980, how long this status would last was uncertain.[11] Each 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. From 1990 onward, the U.S. Congress voted on 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 status contingent on joining the WTO, and this grant came into effect in December 2001 upon China's formal WTO accession. The Trump tariffs of 2018, however, show that gaining PNTR status did not entirely eliminate the risk of further tariff changes.

[8] After 1980, some variation occurs from specific tariffs and temporary commercial policy (as well as aggregation and measurement issues), but this variation is minor relative to the effects of NTR access.

[9] In 1966, President Johnson proposed the East-West Trade Relations Act to provide NTR rates to the Soviet Union and other Warsaw Pact countries. In 1974, the Jackson-Vanik amendment to the Act linked NTR access to freedom of emigration. In 1977, the U.S. International Trade Commission reported on the economic effects of allowing China and the Soviet Union to access U.S. markets at NTR rates.

[10] Due to the One China principle, recognition of the People's Republic of China required withdrawing support for the Republic of China (Taiwan) and the mutual defense treaty. When President Carter did so in December of 1978, debate about its legality was considerable, 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 (Reagan, 1979).

[11] China was the third country to be granted NTR tariffs under the Jackson-Vanik amendment, following Romania in 1975 and Hungary in 1978. After China, no country gained NTR status 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.

2.3 Slow adjustment to the 1980 granting of NTR status

We begin by studying how U.S. imports from China adjusted after the 1980 granting of NTR status. Trade is known to adjust gradually to trade liberalizations. Baier and Bergstrand (2007) find trade doubles in the long run after the creation of a free trade area, but only one-third of the response occurs on impact. Anderson and Yotov (2020), Khan and Khederlarian (2021), and Boehm et 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), 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.

Our first approach is an ECM specification:

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

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 and time-invariant bilateral trade barriers; and good-year fixed effects (δ_{gt}) capture good-level U.S. demand shocks and good-specific trade barriers common to all exporters.[12] Because our sample excludes countries, other than China, that did not receive

[12]This fixed-effects structure has been used frequently in the literature, including by Pierce and Schott

NTR status during the sample, the good-year fixed effects absorb the effects of multilateral changes in NTR tariffs on trade; thus, σ^{SR} and σ^{LR} measure how U.S. imports responded to changes in bilateral tariffs above and beyond 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 in 1979-1981, these elasticities largely capture how Chinese exports responded to the 1980 granting of NTR status.

The solid line in Figure 2 shows the path of adjustment to a one-time tariff change implied by our ECM estimates; the estimated parameter values are reported in the appendix. The short-run trade elasticity is -2.22 and the long-run elasticity is -7.93. The former is consistent with other estimates in the literature, and the latter, although large, is similar to the documented effects of other major liberalizations.[13] The large gap between the short- and long-run responses indicates the adjustment of U.S. imports of Chinese goods to tariff changes has been gradual. Our estimates imply U.S. imports from China take six years to complete 90 percent of the total long-run adjustment to a tariff change. We show in the appendix 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:

$$\Delta_h v_{jg,1979} = \sigma^h_{China} \mathbb{1}_{\{j=China\}} \Delta_h \tau_{jg,1979} + \sigma^h_{Others} \mathbb{1}_{\{j \neq China\}} \Delta_h \tau_{jg,1979} + \delta_{jh} + \delta_{gh} + u_{jg}, \quad (2)$$

where $\Delta_h v_{jg,1979}$ is the h-year log difference in import values relative to 1979: $v_{jg,1979+h} - v_{jg,1979}$, for $h = 1, 2, \ldots, 25$.[14] We follow Boehm et al. (2023) and instrument the h-year

(2016) and Handley and Limão (2017). Anderson and van Wincoop (2003) argue exporter-product-time fixed effects are needed to correctly identify trade elasticities, which in turn requires data on bilateral trade and tariffs for all country pairs. These data are not available over our long sample period. We include these fixed effects in a sensitivity analysis of the NTR-gap elasticity (see section 2.4), and show in the appendix that they do not affect our results.

[13] Khan and Khederlarian (2021) and Alessandria et al. (2024a) 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.

[14] In Alessandria et al. (2024a), 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

change in tariffs relative to 1979 with the tariff change between 1980 and 1979 to account for the autocorrelation of this tariff change.[15] The fixed-effects structure is the same as in (1), except δ_{jg} is eliminated by taking differences of the dependent variable.

The dashed line in Figure 2 shows the path of adjustment to a tariff change implied by our local-projections estimation. Whereas the short- and long-run elasticities obtained by the local projections are very close to those from the ECM, the non-parametric estimation displays a slower transition. Our local-projections estimates imply completing 50 percent of the total adjustment takes 10 years and completing 90 percent takes 19 years. In the appendix, we show these results, too, are robust to a wide range of alternative specifications.

Importantly, these estimates are likely confounded by the effects of uncertainty about the expected persistence of the reform. Our specifications (1) and (2) include only changes in current applied tariffs, 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, the initial response to this reform would have been smaller—and the adjustment process longer—than it would have been in the absence of uncertainty.[16] Thus, we do not interpret these estimates as structural parameters or as causal effects. Instead, we use them to discipline a structural model via indirect inference.

2.4 Effects of the risk of losing NTR status

Our second approach draws from the TPU literature, particularly Pierce and Schott (2016) and Handley and Limão (2017). These studies document that the growth in U.S. imports 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, the United States removed

include only changes relative to 1979.

[15] In the appendix, we show the tariff changes from the NTR access were permanent and slightly increased over time, in contrast to the typical mean-reverting tariff changes over the full sample.

[16] Another potential source of bias is the end of the embargo in 1971. The adjustment to this reform was ongoing when China was granted conditional NTR status in 1980. It also had heterogeneous effects across goods that are correlated with the effects of the 1980 reform. Specifically, goods with lower (higher) NNTR tariffs experienced larger (smaller) liberalizations in 1971 and smaller (larger) liberalizations in 1980.

the annual renewal process, and imports of goods that had faced the largest tariff risk grew fastest. This observatoin 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 back to 1974, before China was granted NTR.

The empirical strategy in the TPU literature is often a difference-in-differences regression that compares goods that were more exposed to future policy risk with goods that were less exposed. In the U.S.-China context, exposure is typically measured by the NTR gap, which is the difference between NNTR and NTR tariffs. We follow Pierce and Schott (2016) and use a time-invariant measure of the NTR gap: $GAP_g = \log(1 + \tau_g^{NNTR} - \tau_{g,2001}^{NTR})$. We include the same set of fixed effects as in section 2.3. Our estimating equation is

$$v_{jgt} = \sum_{t'=1974}^{2007} \beta_t \mathbb{1}_{\{t=t' \wedge j=China\}} GAP_g + \delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}. \tag{3}$$

Our coefficient of interest, β_t, measures how much the exposure to TPU lowered U.S. imports from China in each year relative to imports of the same good in 2008 and relative to imports of the same good from other NTR countries. We refer to this coefficient as the *elasticity of trade to the NTR gap* or the *NTR-gap elasticity*.

Figure 3 shows that between 1974 and 1979, the NTR-gap elasticity was 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 4 shows the NTR gap in 2001 is highly correlated with the changes in applied tariffs on Chinese goods between 1979 and 1981. This correlation also explains why the NTR-gap elasticity prior to 1980 is similar to the long-run trade elasticity estimated in section 2.3.

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 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. The NTR-gap elasticity between the late 1990s and early 2000s is statistically indistinguishable from zero.

In the appendix, we show these results are robust to alternative specifications. Most importantly, we estimate the NTR-gap elasticities using more-aggregated global bilateral imports and exports to consider the role of good-specific Chinese supply factors. This approach controls for spurious correlation between the NTR gap and differences across goods in, for example, export licenses, state-owned enterprises, import quotas, and industry growth.

Just as our estimates of the pace of adjustment to the 1980 granting of NTR status in section 2.3 are confounded by the effects of the risk that this grant could be reversed, our estimates of the effect of this risk are 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 granting of NTR status will diminish in time. Thus, we do not interpret these estimates causally either; instead, they are a means to discipline our model.

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). G goods 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 rate $1 - \delta(z)$, 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, a new firm replaces it exogenously. To export, a firm pays a fixed cost that depends on whether it exported in the previous period. Two trade-policy regimes exist: NNTR and NTR. The probability of switching between regimes varies over time, generating time-varying tariff risk.

Trade policy. Tariffs, $\tau_{gt}(s)$, depend on the current trade-policy regime, s, an aggregate state variable that takes two values: NTR ($s = 1$) or NNTR ($s = 2$).[17] The tariff regime follows a time-varying Markov process with transition probabilities $\omega_t(s, s')$. 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 the appendix.

Production and demand. Firms use labor to produce, $y = z\ell$. Productivity, z, is inde-

[17] We abstract from good-specific risks related to commercial policy.

pendent 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 function of the tariff and the price, p, $d_{gt}(p, s) = (p\tau_{gt}(s))^{-\theta_g} D_{gt}$. D_{gt} is an aggregate demand shifter and θ_g is the good-specific price elasticity of demand. Note the aggregate trade elasticity is determined by the export participation response to a tariff change and the demand elasticity.

Technological trade costs. Firms face two types of technological costs that do not depend on trade policy. One is a stochastic iceberg cost, ξ, that can take three values ($\infty > \xi_{gH} > \xi_{gL}$) and follows 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 nonexporter chooses to export, it begins with ξ_{gH} in the next period. Exporters with $\xi < \infty$ retain their current iceberg costs with probability ρ_ξ and switch with probability $1 - \rho_\xi$. This cost structure implies exporters start exporting small quantities, but with some luck (and repeated investments) grow to export larger quantities. The second technological trade cost is a fixed cost, f, that the firm pays to export in the next period. The fixed costs are identical across firms within a good, but depend on the firm's export history. A nonexporter pays f_{g0} to start exporting the next period. An exporter pays f_{g1} to continue exporting. 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) to capture the exporter life cycle (Ruhl and Willis, 2017). Both types of technological trade costs differ across goods, allowing for heterogeneity in export participation dynamics across sectors in our calibration. The probability of switching variable trade costs (ρ_ξ) is constant across goods because this parameter primarily governs the aggregate long-run trade elasticity.

Firm optimization. Given the firm's export status, it maximizes current-period profits by choosing its price, taking as given its residual demand and the wage, w,

$$\pi_{gt}(z, \xi, s) = \max_p p\, d_{gt}(p, \tau_{gt}(s)) - w \frac{d_{gt}(p, \tau_{gt}(s))\xi}{z}. \tag{4}$$

The values of exporting and not exporting at $t+1$, respectively, are

$$V^1_{gt}\left(z,\xi,s\right) = -f_g(\xi) + \frac{\delta(z)}{1+r}\sum_{s'}\omega_t(s,s')\,\mathop{\mathbb{E}_{\mathrm{t}}}_{z',\xi'} V_{g,t+1}\left(z',\xi',s'\right), \tag{5}$$

$$V^0_{gt}\left(z,\xi,s\right) = \frac{\delta(z)}{1+r}\sum_{s'}\omega_t(s,s')\,\mathop{\mathbb{E}_{\mathrm{t}}}_{z'} V_{t+1}\left(z',\infty,s'\right), \tag{6}$$

where r is the interest rate. The value of a firm is

$$V_{gt}\left(z,\xi,s\right) = \pi_{gt}(z,\xi,s) + \max\left\{V^1_{gt}\left(z,\xi,s\right), V^0_{gt}\left(z,\xi,s\right)\right\}. \tag{7}$$

The break-even exporter has productivity $\bar{z}_{gt}(\xi,s)$, such that $V^1_{gt}\left(\bar{z}_{gt}(\xi,s),\xi,s\right) = V^0_{gt}\left(\bar{z}_{gt}(\xi,s),\xi,s\right)$, which can be rewritten as

$$f_g(\xi) = \frac{\delta(\bar{z}_{gt}(\xi,s))}{1+r}\sum_{s'}\omega_t(s,s')\left\{\mathop{\mathbb{E}_{\mathrm{t}}}_{z',\xi'}\left[V_{t+1}(z',\xi',s')\right] - \mathop{\mathbb{E}_{\mathrm{t}}}_{z'}\left[V_{t+1}(z',\infty,s')\right]\right\}. \tag{8}$$

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. 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

$$\varphi_{g,t+1}(\mathcal{Z},\infty) = \sum_{\xi}\left[\int_0^{\bar{z}_{gt}(\xi,s_t)} Q_{gt}(\mathcal{Z},z,\xi)dz + \int_{\bar{z}_{gt}(\xi,s_t)}^{\infty}\bar{h}_g(\mathcal{Z})\varphi_{gt}(z,\xi)dz\right], \tag{9}$$

$$\begin{aligned}\varphi_{g,t+1}(\mathcal{Z},\xi_{gH}) = {} & \int_{\bar{z}_{gt}(\infty,s_t)}^{\infty} Q_{gt}(\mathcal{Z},z,\infty)dz + \rho_\xi\int_{\bar{z}_{gt}(\xi_{gH},s_t)}^{\infty} Q_{gt}(\mathcal{Z},z,\xi_{gH})dz \\ & + (1-\rho_\xi)\int_{\bar{z}_{gt}(\xi_{gL},s_t)}^{\infty} Q_{gt}(\mathcal{Z},z,\xi_{gL})dz,\end{aligned} \tag{10}$$

$$\varphi_{g,t+1}(\mathcal{Z},\xi_{gL}) = (1-\rho_\xi)\int_{\bar{z}_{gt}(\xi_{gH},s_t)}^{\infty} Q_{gt}(\mathcal{Z},z,\xi_{gH})dz + \rho_\xi\int_{\bar{z}_{gt}(\xi_{gL},s_t)}^{\infty} Q_{gt}(\mathcal{Z},z,\xi_{gL})dz, \tag{11}$$

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, $\bar{h}_g(\mathcal{Z})$ is the probability of dying and being replaced by a new firm with productivity in $\mathcal{Z}$, and $Q_{gt}(\mathcal{Z}, z, \xi) = h_g(\mathcal{Z}, z)\varphi_{gt}(z, \xi)$. Although the decision rules, $\bar{z}_{gt}(\xi, s)$, respond immediately to trade-policy changes, the stock of exporters across trade costs adjusts gradually. Consequently, aggregate trade volumes,

$$EX_{gt}(s) = \sum_{\xi \in \left\{\xi_{gL}, \xi_{gH}\right\}} \int_z p\left(z, \xi, \tau_{gt}(s)\right) y\left(z, \xi, \tau_{gt}(s)\right) \varphi_{gt}\left(z, \xi\right) dz, \tag{12}$$

respond slowly to policy changes. The slow adjustment 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 has four stages. First, we group our 5-digit SITC goods into 15 sectors and use Chinese firm-level data to compute sector-level statistics about exporter dynamics from 2004–2007. Second, we assign several parameters to standard values from the literature and make some common functional-form assumptions. Third, we calibrate the sector-specific parameters so that the model's terminal steady state matches the moments from the first stage. 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. The calibration is summarized in Table 3.

4.1 Chinese firm-level data and sectoral heterogeneity

Following the trade-dynamics literature, we calibrate the 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.[18] We compute four statistics that summarize the distribution and dynamics of Chinese firms that export to the United States: the dispersion in export sales (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 incumbent exporters divided by the average exports of new exporters (the incumbent premium).

Aggregation bias is a potential concern. Exporter dynamics could vary across sectors, reflecting sectoral heterogeneity in non-tariff trade costs and other technological primitives. If this heterogeneity were correlated with the NTR gap, the evolution of NTR-gap elasticity (Figure 3) 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 toward China.[19] To ensure our results are not driven by aggregation bias, we compute statistics for each sector in the 2-digit China Industry Classification System (Table 1). We map the 5-digit SITC goods in our data to this system using the ISIC Revision 4 concordance. Heterogeneity exists 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. We find a positive relationship between the NTR gap and the export participation rate (correlation 0.51), a negative relationship with the exit rate (–0.55), and no clear relationship with the coefficient of variation of log exports (0.24), and the incumbent premium (0.03).

4.2 Assigned parameters and functional forms

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 NTR 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

[18] These data are widely used to study Chinese manufacturing growth in the late 1990s and 2000s (see Bai et al., 2023). We thank Dan Lu for sharing the data. See the appendix for more details about these data.

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

Alessandria et al. (2021). Firm productivity follows $\log a' = \rho_z \ln a + \varepsilon$, with ε normally distributed with mean zero and standard deviation σ_{gz}, and $z = \frac{1}{\theta_g - 1} \log a$. This specification eliminates the role of the demand elasticity, θ_g, in the size distribution of firms, which facilitates computation. We assume the persistence parameter, ρ_z, is common across goods, while the variance of the innovations, σ_{gz}, varies.[20] 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}.

We use estimates of U.S. import demand elasticities in Soderbery (2018) to set the analogous elasticities in our model, θ_g.[21] θ_g is the same for all goods within a sector but varies across sectors. We concord Soderbery's (2018) 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 15 sectors by taking the average. Table 2 reports our sectoral demand elasticities, which range from 2.7 to 3.7. No systematic relationship exists 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 in the 2004–2007 data (Table 1). Our empirical results show 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 perspective 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. We have $15 \times 4 = 60$

[20] The Chinese firm-level data are too short 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 is cleanly identified by cross-sectional variation in export sales.

[21] We use Soderbery's (2018) 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.

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 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 2 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. We target the long-run trade elasticity of -7.93 from the error-correction model and the annual NTR-gap elasticities shown in Figure 3.[22] 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 1971–1979 and NTR from 1980 onward), update the distributions $\varphi_{g,t}$ using the laws of motion (9)–(11), and compute aggregate exports using (12). We then estimate equations (1) and (3) on the simulated data.

The parameters calibrated in this stage are exactly identified: the number of parameters is the same as the number of target coefficients. The probability of switching trade costs, ρ_ξ, is primarily identified by the long-run aggregate trade elasticity.[23] This parameter governs the measure of high-capacity exporters (firms with $\xi = \xi_{gL}$) in the long run and 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 NTR, $\omega_t(2, 1)$, is identified by the NTR-gap

[22]We HP-filter the NTR-gap coefficients from 1981 onward to smooth out temporary spikes (e.g., 1984). The calibrated model exactly matches this smoothed series, shown in Figure 5.

[23]Alternatively, we could target the long-run trade elasticity from the local-projections specification (2), but these elasticities are very similar, so the implied parameter values would be very similar.

elasticity during the 1970s, when China was in the NNTR regime. A higher probability of gaining NTR status boosts exports more in high-gap industries than in low-gap industries because the former gain more from this status. Given the relatively short time China is in the NNTR regime, and the relatively flat NTR-gap elasticity during this period, we assume $\omega_t(2,1)$ is constant and target the average NTR-gap elasticity in 1974–1979.

The probability of switching from NTR to NNTR, $\omega_t(1,2)$, is identified by the NTR-gap elasticity from 1981 onward. The probability of losing NTR 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 NTR status reduces export participation more in high-gap industries than in low-gap industries, but this finding does not affect trade volumes until the following period.[24]

Note this stage of the calibration procedure is an indirect-inference exercise in the spirit of Gourieroux et al. (1993). The ECM and NTR-gap specifications, (1) and (3), are misspecified: the former is confounded by changes in expectations about future policy, whereas the latter is confounded by gradual adjustments to past changes in applied tariffs. These issues apply to the model as well. Indirect inference allows us to use the estimates from these specifications to recover our model's structural parameters.

5 Results

First, we discuss how trade-policy expectations have evolved since 1980. Second, we study the contribution of TPU 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. Last, 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.

[24]See the appendix for more detail on identification of these probabilities and the sensitivity of our estimates to the NTR-gap elasticity path.

5.1 Estimates of trade-policy expectations

We begin with the main result of our calibration: the annual probability of switching between policy regimes (Figure 5). The probability of switching from the NNTR to the NTR regime was about 28 percent. The probability of switching back to the NNTR regime was initially 19 percent in 1980, rose sharply to 59 percent in 1981, and fell throughout the mid 1980s and early 1990s. A temporary increase occurred in 1994–1996 and a smaller increase in the early 2000s, but the overall trend continued downward. By 2008, the end of our observation period, the probability of moving back to the NNTR regime had fallen to 1.1 percent.[25]

The historical context supports our estimates of trade-policy expectations in the years surrounding the 1980 reform.[26] 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 many additional steps to NTR still remained 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

[25] That the probability does not fall to zero implies trade is permanently depressed, albeit slightly, by the possibility of losing NTR status. This positive probability is consistent with evidence from the U.S.-China trade war, where trade in high-NTR-gap goods rose relative to trade in low-NTR-gap goods when new tariffs on Chinese goods were introduced in late 2018. These tariffs were 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. We study policy uncertainty during the trade war in Alessandria et al. (2024b).

[26] Much of this discussion is based on a series of reports that provide summaries of key changes in trade and trade policy (USITC, 1970–1990). For additional reading, see the appendix.

losing NTR status from 1985–1993 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 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 small change in the probability of moving back to the NNTR regime when China gained PNTR status in 2001. The probability was about the same in the mid 2000s as in the early 1990s. This finding tells a different story about the effect of gaining PNTR than Pierce and Schott (2016) and Handley and Limão (2017), who argue the change in the NTR-gap elasticity after 2001 is evidence that PNTR significantly reduced the probability of losing NTR status. Two key factors explain this difference. First, in section 5.2, we show gradual adjustment to the 1980 granting of NTR status (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 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 each year. Figure 5 plots the discounted expected tariff across goods in the model,

$$\tau_t^{PV} = \frac{1}{G}\sum_{g=1}^{G}(1-\beta)\left(\sum_{s=t}^{\infty}\beta^{s-t}\mathbb{E}_t[\tau_{gs}]\right), \tag{13}$$

and the mean applied tariff. Whereas the realized path of applied tariffs fell sharply in 1980, and then fell slightly throughout the 1980s and 1990s from reforms to U.S. NTR tariff rates, the discounted expected tariff fell gradually throughout the entire period. The discounted expected tariff exceeded the applied NTR 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. 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, 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 NTR regime, but thereafter, they believe the NTR regime will last forever. 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.[27]

Figure 5 plots the NTR-gap elasticities obtained by estimating (3) 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 NTR 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

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

actual elasticity at each point in time measures the contribution of policy uncertainty. This contribution was, on average, 61 percent of the elasticity in the data during the 1980s, 56 percent during the 1990s, and 73 percent during the 2000s. Note 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). The reason is that delayed adjustment to the 1980 reform played 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 TPU on aggregate trade. Figure 5 plots aggregate exports in the counterfactual and the benchmark models. Again, the vertical distance between the two lines measures the effect of TPU. 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, which permanently reduces the number of exporters in high-gap industries. This finding is consistent with Alessandria et al. (2024b), who find 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 NTR 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 granting of NTR status (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 TPU 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):

$$v_{jgt} = \beta \mathbb{1}_{\{t<2000 \wedge j=China\}} X_g + \sigma \tau_{jgt} + \delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}. \tag{14}$$

When we estimate this regression using the observed data, we find $\beta = -0.67$, and we obtain essentially the same estimate using simulated data from our benchmark model with policy uncertainty.[28] When we estimate this regression on simulated data from the no-TPU counterfactual, we find $\beta = -0.30$, which is smaller than the estimate from the data but still statistically significant. This finding shows that even if the permanence of China's NTR status had never been in doubt, exports of high-gap goods still would have grown faster than imports of low-gap goods after China gained PNTR status in 2001. Quantitatively, it indicates a gradual adjustment to the earlier reform accounts for about 40% of the overall effect of PNTR on trade as documented by Pierce and Schott (2016).

5.4 The role of time-varying policy uncertainty

We find that changes in expectations about U.S. trade policy toward China mostly occurred in the late 1980s and early 1990s, rather than, as others have argued, at the turn of the century. Here, we study how U.S. imports from China would have grown if only one or two changes in expectations occurred at key geopolitical moments.

First, we study 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.[29] 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 (14). Figure 6 shows this version of the model (*Const. TPU from 1980*) does fairly well in matching the NTR-gap elasticity during the 1990s and 2000s. Trade grows too quickly, however, in high-gap industries relative to

[28] Our estimate is slightly different than Pierce and Schott's (2016) because we use a different level of aggregation. See the appendix for additional details.

[29] 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 3.

low-gap industries in the 1980s. The probability of losing NTR status in this version of the model is about five percent, which is more than twice the decline in the average probability in the benchmark model between 1992–2000 and 2001–2007.

Next, we consider a model with no risk of losing NTR status prior to 1990 but a constant risk in the 1990s. This model is intended to capture the idea that non-renewal was not a serious concern until Congress began voting annually on China's NTR status in 1990—a common assumption in the TPU literature.[30] Figure 6 shows this model (*Const. TPU from 1990*) is inconsistent with the data during all time periods. The NTR-gap elasticity in this model shrinks rapidly during the early 1980s—it follows the no-TPU counterfactual's trajectory exactly—before growing again during the 1990s when the risk of losing NTR status arises, and then shrinking rapidly again following China's 2001 WTO accession. The probability of losing NTR status in this model is 16 percent, which is well above the benchmark model's highest probability in the 1990s and more than 8 times the difference between that model's 1992–2000 and 2001–2007 averages.

These analyses show China's export growth cannot be understood without time-varying policy uncertainty, particularly in the early years after NTR status was granted. The 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 from China over the next two decades, and ignoring this trend overstates the degree to which uncertainty fell after China gained PNTR status in 2001. The analyses 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 and past changes in tariffs.

[30] 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).

6 Employment effects

Changes in policy uncertainty played the largest role in export growth during 1986–1993. At first glance, this narrative is inconsistent with Pierce and Schott (2016), who document a large decline in U.S. employment in high-gap industries after China joins the WTO. We revisit these employment effects over our longer period. Once we introduce the industry controls suggested by theory, the decline in employment from trade policy on Chinese imports starts 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 is 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 on China.

6.1 Conceptual framework

To set ideas, consider an Armington model of U.S. demand[31] with an aggregate consumption good, Q_t, that is a bundle of industry-level goods, Q_{gt}, with an elasticity of substitution α,

$$Q_t = \left(\sum_g Q_{gt}^{\frac{\alpha-1}{\alpha}} \right)^{\frac{\alpha}{\alpha-1}}. \tag{15}$$

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,

$$Q_{gt} = \left(Q_{Dgt}^{\frac{\theta_g-1}{\theta_g}} + Q_{Mgt}^{\frac{\theta_g-1}{\theta_g}} \right)^{\frac{\theta_g}{\theta_g-1}}. \tag{16}$$

[31]This demand system is consistent with GE analyses of Chinese integration by Kehoe et al. (2018), Caliendo et al. (2019), and Galle et al. (2022). These papers consider a broader set of shocks than changes in trade from trade policy. It is also consistent with the demand function in the model in section 3.

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

$$P_{Dgt}Q_{Dgt} = \left(\frac{P_{Dgt}}{P_{gt}}\right)^{1-\theta_g} P_{gt}Q_{gt}, \tag{17}$$

$$P_{gt} = \left(P_{Dgt}^{1-\theta_g} + (P_{Mgt}\tau_{Mgt})^{1-\theta_g}\right)^{\frac{1}{1-\theta_g}}, \tag{18}$$

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 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 are $P_{Dgt}Y_{gt} = P_{Dgt}Q_{Dgt} + P_{Dgt}Q^*_{Mgt}$, where Q^*_{Mgt} is the quantity demanded in China and we have assumed the export price and domestic price are identical.

Production in U.S. industry g is $Y_{gt} = Z_{gt}L_{gt}$, where Z_{gt} is labor productivity. Combining (17), (18), and the definition of $P_{Dgt}Y_{gt}$, and log-linearizing, we arrive at

$$\begin{aligned} d\ln L_{gt} \approx &(1-\omega_{Xg})[\omega_{Mg}(\theta_g - 1)(d\ln P_{Mgt}\tau_{Mgt} - d\ln P_{Dgt}) \\ &+ d\ln(P_{gt}Q_{gt}) - d\ln P_{Dgt}] + \omega_{Xg}\left[d\ln(P^*_{Mgt}Q^*_{Mgt}) - d\ln P_{Dgt}\right] - d\ln Z_{gt}. \end{aligned} \tag{19}$$

This equation shows 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. The industry's exposure to imports and exports, as captured by ω_{Xg} and ω_{Mg}, amplify or attenuate these effects. 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.

6.2 Estimation details

Relative to our work with U.S. import data, two challenges arise in estimating (19). 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). 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 the prospect of import growth in the far future was unlikely to be a material factor.

The second challenge is the lack of data on the relative price of Chinese imports to locally produced goods, $d \ln P_{Mjt}\tau_{Mjt} - d \ln P_{Djt}$. The literature, however, focuses 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 a 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 (3):

$$\begin{aligned} d \ln L_{gt} =& (1-\omega_{Xg}) \left[\omega_{Mg} \sum_{t'=1958}^{2007} \beta_t \mathbb{1}_{\{t=t'\}} GAP_g + d\ln(P_{gt}Q_{gt}) - d\ln P_{Dgt} \right] \\ &+ \omega_{Xg}(d\ln(P^*_{Mgt}Q^*_{Mgt}) - d\ln P_{Dgt}) - d\ln Z_{gt} + \delta_g + \delta_t + \epsilon_{gt}, \end{aligned} \tag{20}$$

where we have further included industry (δ_g) and time (δ_t) fixed effects.[32] Note the shares (ω_{Xg}, ω_{Mg}) are time invariant. These shares are measured in the base period around which we linearize. We use the average shares over 1995–1999 for ω_{Xg} and ω_{Mg}.[33] Standard errors are clustered at the 4-digit SIC industry level.

6.3 Estimation results

Our baseline specification and alternatives are in Figure 7. We estimate one set of parameters for each industry and report results for the industry in the 90th percentile of Chinese import share and the 50th percentile export share.[34] Plotted in Figure 7 is $\beta_t(1-\omega_{Xg})\omega_{Mg}$, the

[32]By imposing the same elasticity on all industries, this specification ignores heterogeneity in θ_g. Because θ_g is orthogonal to the NTR gap, this heterogeneity should not affect our estimate of β_t.

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

[34]In the appendix, we consider 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

elasticity of employment to the NTR gap. Employment falls more in industries that are more exposed to the NTR gap, but most of the decline occurs before China joins the WTO.

To compare our estimates with Pierce and Schott (2016), we estimate

$$d \ln L_{gt} = \sum_{t'=1958}^{2007} \beta_t \mathbb{1}_{\{t=t'\}} GAP_g + \delta_g + \delta_t + \epsilon_{gt}. \tag{21}$$

Relative to (20), this specification forces the employment elasticity to be constant across industries and drops the industry controls for domestic expenditures, exports, and TFP. Figure 7 shows that, without these controls, employment dynamics are very different: employment is stable during 1990–1997, 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 did not trade with the U.S. then. The large differences in moving from the theoretical regression in (20) to the reduced-form regression in (21) suggests the decline in employment in high-gap industries around the time that China joined the WTO may not reflect a change in TPU.

We gain some insight into the importance of the theory-motivated controls by considering alternative versions of (20) in Figure 7. Removing TFP changes our results a 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, TFP, and export and domestic demand, leaves us with (21), 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, reconciles our findings with those of Pierce and Schott (2016).

The difference between our specification and (21) can be further understood by studying the components of the employment elasticity, $\beta_t(1-\omega_{Xg})\omega_{Mg}$. Our results in Figure 7 change the most when we do not control for heterogeneous import and export exposure through ω_{Xg}

effects with ω_M. Given the much smaller variation in domestic sales shares, less variation is present in the employment effects with ω_X.

and ω_{Mg}. As we show in the appendix, industries with higher Chinese import shares tend to have larger NTR gaps, rather than smaller gaps as one might expect, which suggests China has a comparative advantage in these industries. Allowing the elasticity of employment to the NTR gap to vary with import and export exposure is important, lest we attribute employment declines in high-gap industries to multilateral China supply factors.

The NTR-gap elasticities, β_t, estimated with employment data should, in theory, be the same as the elasticities (with the opposite sign) estimated with the import data from (3). Absent issues of aggregation, the main differences are (i) (20) lacks the rich controls available with the import data, and (ii) the substitution identified is between imports and domestic sales, rather than across source countries. To gauge the importance of these considerations, in Figure 7, we plot $\beta_t(1-\omega_{Xg})\omega_{Mg}$ where we use the estimates of β_t from the import regression (3) ("trade coefficients" in the figure) instead of the estimates from the employment regression (20). The effects on employment are half as big as those in our baseline results using employment data.

7 Conclusions

We study, empirically and quantitatively, the growth of China's exports to the United States since the 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 the initial lack of credibility about the 1980 granting of NTR status 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 trade-policy expectations leverages unique aspects of U.S. policy toward China in which potential changes in future trade policy are known and heterogeneous across products, whereas the likelihood of this change is unknown and common across products. Our analysis could be extended to consider other 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 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.

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. (2024a) 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.

Data availability

Code replicating the tables and figures in this article can be found in Alessandria et al. (20024) in the Harvard Dataverse, `https://doi.org/10.7910/DVN/GSH8OM`.

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Table 1: Chinese exporter dynamics statistics, 2004–2007

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

Notes: The data are described in the appendix. Reported moments are averages for 2004–2007.

Table 2: Sector-level model parameters

	Sector	θ_g	f_{g0}	f_{g1}	ξ_{gH}	σ_{gz}
1	Food, beverage, tobacco	3.09	0.14	0.12	3.34	0.91
2	Textile, clothing, footwear	3.17	0.20	0.13	2.57	1.02
3	Wood and straw products	2.79	0.26	0.17	3.71	1.03
4	Paper, printing products	3.43	0.19	0.16	3.46	1.08
5	Energy products, chemicals	2.99	0.27	0.20	4.56	1.17
6	Rubber, plastic products	3.16	0.18	0.12	3.40	0.99
7	Non-metallic mineral products	2.85	0.16	0.14	3.56	0.90
8	Base metal manuf.	3.04	0.13	0.15	4.60	0.99
9	Calendered metal manuf.	2.73	0.29	0.17	4.62	1.07
10	Other machinery, equipment	3.74	0.23	0.15	3.03	1.20
11	Computer, electronic, optical	3.18	0.46	0.20	4.81	1.35
12	Electrical equipment manuf.	3.27	0.30	0.16	3.86	1.20
13	Vehicle manuf.	3.06	0.20	0.15	4.81	1.07
14	Furniture, other manuf.	3.26	0.20	0.11	2.25	0.98
15	Non-manufacturing	2.97	0.22	0.16	4.04	1.07

Table 3: Calibration summary

Parameter	Meaning	Value	Source/target
(a) Assigned			
r	Interest rate	4 pct.	Standard
ρ_z	Persistence of prod.	0.65	Alessandria et al. (2021)
δ_0	Corr.(survival, prod.)	21.04	Alessandria et al. (2021)
δ_1	Minimum death prob.	0.023	Alessandria et al. (2021)
τ_{g1}	NNTR tariff	Varies	Data
τ_{g2}	NTR 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}	Prod. dispersion	Varies	CV of log sales
(c) Calibrated to transition			
ρ_ξ	Prob. of keeping iceberg cost	0.91	LR trade elast. $= -7.93$
$\omega(2,1)$	Prob. NNTR to NTR	0.28	NTR-gap elast., 1974–79
$\omega_t(1,2)$	Prob. NTR to NNTR	Varies	NTR-gap elast., 1981–08

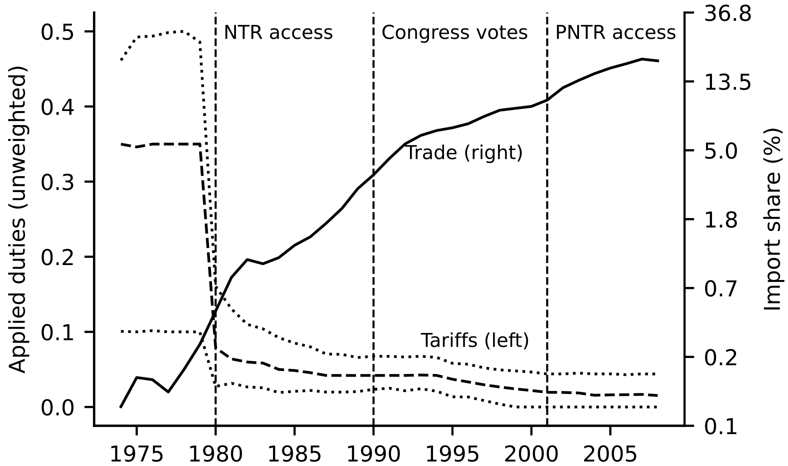

Fig. 1 – Growth of aggregate imports from China. Solid line: China's log share of total U.S. imports. Dashed line: median tariff across goods. Dotted lines: 25th and 75th percentiles.

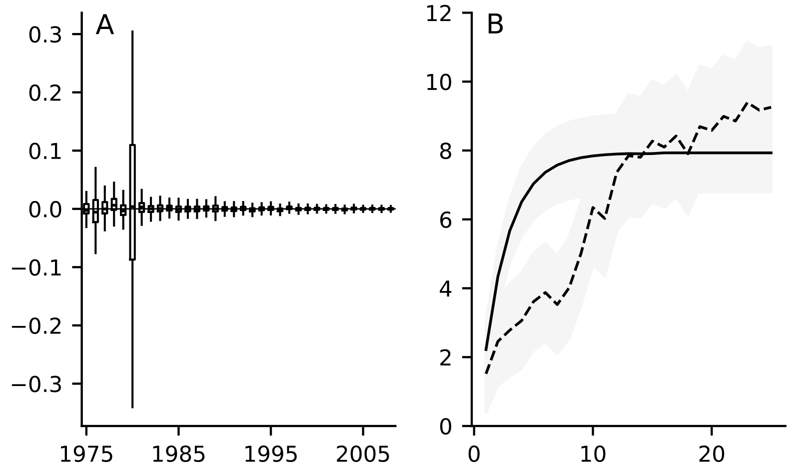

Fig. 2 – A: box-and-whisker plot of distribution of residuals from regressing $\Delta_1 \tau_{jgt}$ on δ_{jt}, δ_{jg}, and δ_{gt}. B: elasticities of U.S. imports from China to tariff changes. Solid line shows ECM estimates from (1). Dashed line shows local-projections estimates from (2). Shaded areas: 95-percent confidence intervals.

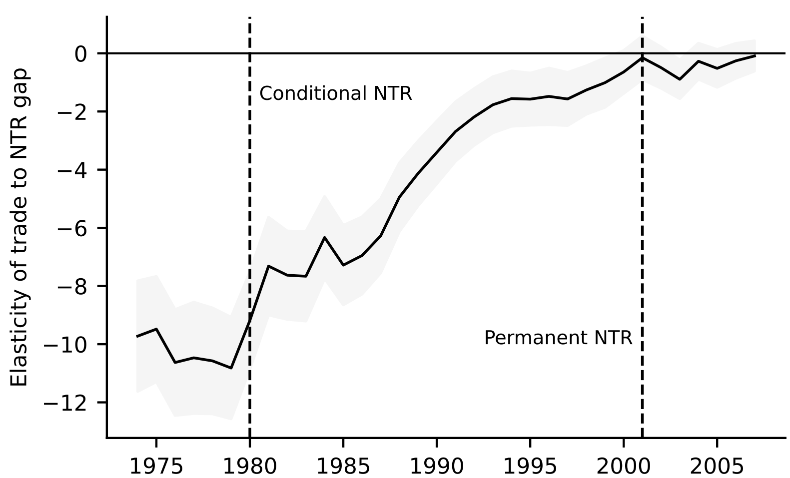

Fig. 3 – Elasticity of U.S. imports from China to the NTR gap. Solid line: estimates of $\hat{\beta}_t$ from (3). Shaded area: 95-percent confidence interval.

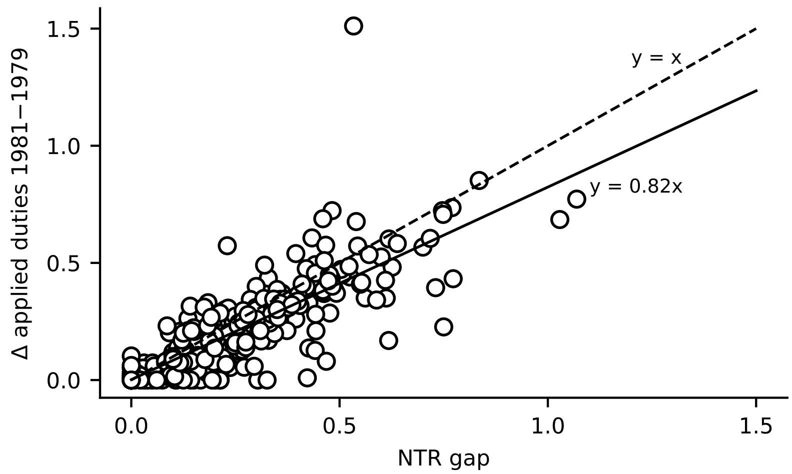

Fig. 4 – Size of 1980 liberalization vs. NTR gap. Each dot is one 5-digit SITC good. Line of best fit has an R-squared of 0.86.

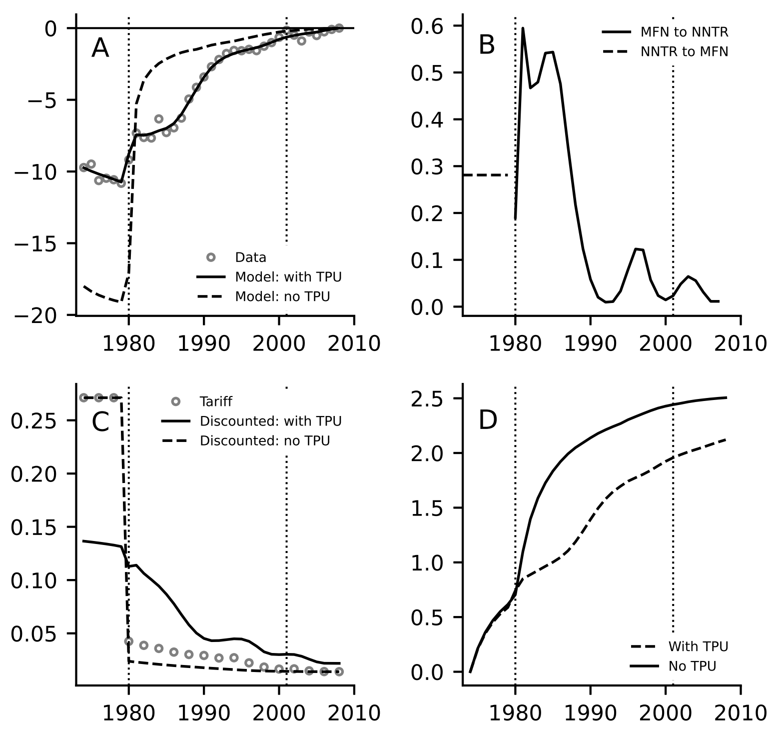

Fig. 5 – Model results. A: NTR-gap elasticities. B: policy transition probabilities. C: discounted expected value of tariffs. D: aggregate trade.

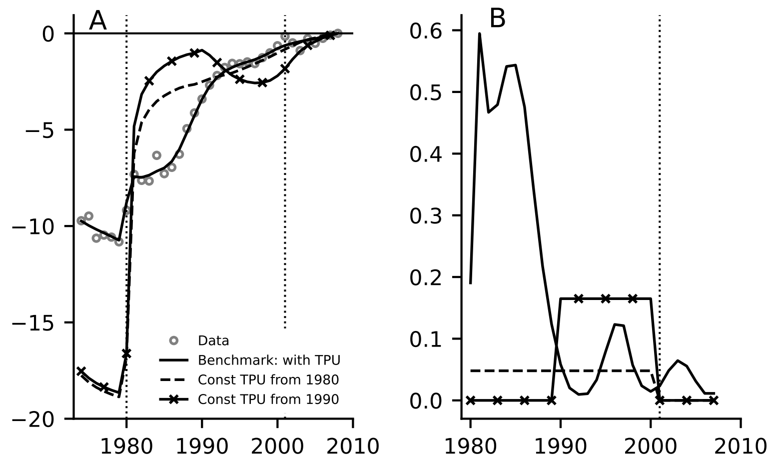

Fig. 6 – Benchmark model vs. constant-TPU models. A: NTR-gap elasticities. B: policy transition probabilities.

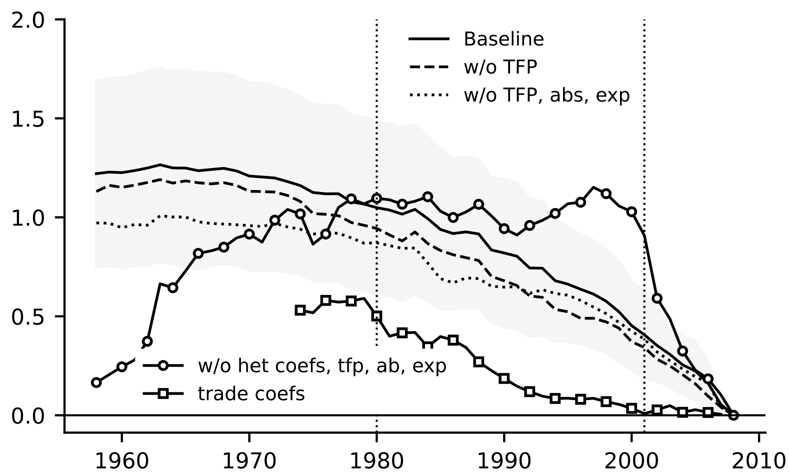

Fig. 7 – Elasticity of U.S. employment to NTR gap. Solid line: estimates from (20) for industry with median domestic sales share and 90$^{\text{th}}$ percentile Chinese import share. Shaded area: 95-percent confidence interval. Dashed line: drop TFP control. Dotted line: drop all controls. Circles: drop all controls and remove interactions with industry shares.

Appendix to: Trade-policy dynamics: Evidence from 60 years of U.S.-China trade

Alessandria, Khan, Khederlarian, Ruhl, and Steinberg

July 2024

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 G and H contain the additional tables and figures discussed in the aforementioned sections.

Contents

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.[1] 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 G.2, 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.[2]

B Robustness: Slow adjustment

The results of the ECM and LP specifications presented Section 2.3 are robust to a number of alternative specifications. For visualization purposes we only present the results corresponding to the ECM. Results of the corresponding local projection results are available upon request.

Shipping costs. Column 2 of Table G.3 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 G.3 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.

[1] 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.

[2] We obtain this concordance from Marc Muendler's website.

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 G.4 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 3 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 (3). 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 G.5 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 $\hat{\beta} = -0.9$.[3] 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 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.

[3] 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.

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 (3) of the paper. Figure H.2 shows that the extensive margin clearly plays an important role in the evolution of the overall NTR gap.[4]

China supply effects. Our baseline approach controls for good-specific U.S. demand shocks but not for good-specific Chinese supply shocks, such as export licensing, privatization of

[4] 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 H.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 H.2 as a quantitative decomposition of the contributions of the extensive and intensive margins to the overall NTR-gap elasticity.

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).[5] 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 H.4. The solid line plots the response of U.S. imports from China when considering U.S. imports only. This is analogous to (3), 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

[5] 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.

determined, namely the TSUSA product schedule between 1974–1988 and the HS product 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 H.5 plots the results of (3) 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.[6] Second, we estimate (3) for the continuous sample period using a new concordance from Acosta and Cox (2022) between the TSUSA and the HS schedules.[7] The NTR-gap elasticities under this approach are reported in Figure H.6. Again, the path of the annual NTR-gap elasticities is similar.

Sample of countries and goods. Table G.8 demonstrates that our estimates of (3) 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).[8] 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 (3) we consider the 1999

[6] We also report robustness to these results in Table G.6 (TSUSA) and Table G.7 (HS-8).

[7] 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.

[8] 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.

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 joined the WTO. To illustrate that most of the variation is due to the NNTR rate we have estimated (3) defining Gap_g as the NNTR rate only. The results are reported in Column 1 of Table G.9. 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 4) we have also estimated (3) 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 G.9). 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 G.9.

Additional trade-cost controls. Our baseline estimation (3) 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 4). Columns 4–5 of Table G.9 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 3, 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 (3) 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 G.9 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 seven 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 the sensitivity of our results to the parameters that govern exporter life-cycle dynamics, which play important roles in determining the long-run effects of trade reforms and the length of transitions. In the third, we study a version of the model in which there is no parameter heterogeneity across sectors. In the fourth, we study a version of the model without exporter life-cycle dynamics at all to further illustrate the role of gradual adjustments. In the fifth, 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 sixth, we provide some additional results that illustrate how the trade-policy transition probabilities are identified from the NTR-gap elasticities. 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 3, 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.[9] The results are shown in Figure H.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 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.

Sensitivity to exporter life-cycle parameters. We have estimated the parameters that govern exporter life-cycle dynamics—the high iceberg trade cost that applies to new entrants, ξ_{gH}, and the probability of switching to the low iceberg cost, $\rho_x i$—to match empirical facts that we have estimated. These parameters are particularly important because they play crucial roles in determining how much trade responds to changes in policy in the long run and how long these responses take to materialize in the absence of policy uncertainty. To study the sensitivity of our results to these parameters, we perturb each of these parameters and re-estimate the trade-policy transition probabilities, $\omega_t(s, s')$.

Figure H.8 shows our results in alternative calibrations in which we use different values of

[9] 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.

the high iceberg cost, ξ_{gH}, while Figure H.9 shows the results in alternative calibrations with different values of the iceberg cost transition probability, ρ_ξ. Panel (a) in each of these figures shows that in the absence of trade-policy uncertainty, the NTR-gap elasticity would follow a materially different path than in the benchmark calibration, both in terms of the level in the 1970s (which is closely related to the long-run effect of trade policy) and the speed of adjustment. Nevertheless, panel (b) shows that the estimated trade-policy probabilities in these alternative calibrations are very similar to the benchmark probabilities. Qualitatively, our main results—that the likelihood of losing NTR status was initially very high in the early 1980s, fell dramatically during the late 1980s, and did not change materially when China joined the WTO in 2001—hold up in all of these alternative calibrations.

The role of realistic exporter life-cycle dynamics. A different way to study the role of gradual adjustment than the accounting exercise we conducted in Section 5.3 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 H.11, 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 H.11). 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.

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 H.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.

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 H.12 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.

Identification of policy transition probabilities from NTR-gap elasticities. Here we provide some additional results that illustrate how the trade-policy transition probabilities are identified from the NTR-gap elasticities. Recall that the probability of gaining NTR status during the 1970s, $\omega(2,1)$, is identified from the average NTR-gap elasticity during this period, and that the probability of losing NTR status in period t, $\omega_t(1,2)$, is identified by the NTR-gap elasticity in period $t+1$.

To show how this identification works in practice, Figure H.13 shows how the NTR-gap elasticity path in the model changes when we increase each probability by 1 percentage point, one probability at a time. In each panel of the figure, the elasticity used to identify the probability in question is shown in red, while the other elasticities are shown in blue. Y-axis scales are omitted to emphasize which elasticities are relatively more affected by each probability, rather than the magnitude of these effects. The first panel of the figure shows how the NTR-gap elasticities change when the probability of gaining NTR status, $\omega(2,1)$, increases. This shifts the entire elasticity path upward, but the elasticities during the 1970s, which are used to identify this probability, change the most. The remaining panels show how the elasticities change when the probability of losing NTR status in a given year, $\omega_t(1,2)$, increases. In each case, although the entire elasticity path shifts downward, the effects are local, in the sense that the elasticity in period $t+1$ used to identify this probability shifts the most, and elasticities further back or further forward shift less. This illustrates the essence of how our identification strategy works.

To get a clearer sense of the identification, it is helpful to conduct the same exercise in the *surprises* model studied above. In this model, the probability of losing NTR status in period t only affects the NTR-gap elasticities in period $t+1$ onward; it cannot affect the elasticities in prior years because the year-to-year transition probability changes are unanticipated. The results are shown in Figure H.14. One can clearly see now how the identification proceeds sequentially. First, the probability of losing NTR status in 1980, $\omega_{1980}(1,2)$, is identified from the NTR-gap elasticity in 1981, because the probabilities from 1981 onward have no effect on this moment. Then, having pinned this parameter down, we can use the NTR-gap elasticity in 1982 to identify $\omega_{1981}(1,2)$, because the probabilities from 1982 onward do not affect this moment. We proceed in this way forward in time until finishing the identification by pinning down $\omega_{2006}(1,2)$ from the NTR-gap elasticity in 2007. Since the estimated probabilities in this version of the model are very similar to our benchmark estimates, the economic intuition above still largely holds in the benchmark model, even though it does not hold exactly. This is confirmed by the "locality" of the effects of the probabilities on elasticities close in time to the moments used for identification discussed above.

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 $b = 1, 2, \ldots, 5$, we set a so that the mean prior, given by $a/(a+b)$, equals 0.72. The prior with $b = 1$ represents an agent with little confidence in this value, whereas the prior with $b = 5$ represents a highly confident agent. Panel (a) of Figure H.15 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 H.19(a) shows a minor break in the gap elasticity around the switch periods.

Figure H.16 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. As we write in the main text, 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.

Figure H.17 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 H.18 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.

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 (20) and our reduced-form specification (21). Figure H.19 plots the elasticity of several key variables to the NTR gap from 1958–2018. Figure H.19(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 H.19(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.

Figure H.20 compares our estimates of β_t with trade and employment data. With employment data, the NTR-gap elasticities are about twice as large as those from the import data. The employment data shows that the substitution is much later than in the trade data. To reconcile these findings, we remove several of the fixed effects from our trade regression that are omitted from our employment regression. We estimate (3) with only time fixed effects and product-country fixed effects. With these minimal fixed effects, import substitution towards high-gap goods that is almost twice as large as the 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.

F U.S. trade policy towards China

The discussion in section 5.1 is based on a series of reports that provide summaries of key changes in trade and trade policy (USITC, 1970–1990). For further background reading, see: 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."

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G Additional tables

Table G.1: Summary statistics for NNTR and NTR tariff schedules

						Applied Duties			
		NNTR Rate		U.S. Export Share		1979		2001	
	SITC 1-Digit	Mean	Std.	1979	2001	China	NTR	China	NTR
0	Food and live animals	16	14	13	1	15	4	3	2
1	Beverages and tobacco	48	45	0	0	49	14	3	2
2	Crude materials, inedible, except fuels	11	17	17	1	6	1	0	0
3	Mineral fuels, lubricants and related	1	1	25	0	1	0	1	0
4	Animal and vegetable oils, fats and waxes	11	7	1	0	4	2	2	2
5	Chemical and related products	29	15	14	2	20	5	3	2
6	Manufactured goods	36	17	12	10	38	7	2	2
7	Machinery and transport equip.	34	11	0	40	33	5	1	1
8	Misc. manufactured articles	39	23	17	45	40	7	3	2
9	Commodities and transactions, n.e.c.	33	24	1	0	28	3	1	1
10	Average	29	21			28	6	2	2

Table G.2: 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 G.3: Slow adjustment with ECM — robustness

	Baseline	Shipping	All Goods	All Countries	Full	Balanced
$\mathbb{1}\{j \neq China\}\Delta\tau_{jgt}$	-2.06***	-2.06***	-1.90***	-2.24***	-2.40***	-1.95***
	(0.14)	(0.14)	(0.11)	(0.13)	(0.10)	(0.15)
$\mathbb{1}\{j = China\}\Delta\tau_{jgt}$	-2.44***	-2.44***	-2.63***	-2.48***	-2.89***	-2.44***
	(0.40)	(0.40)	(0.41)	(0.40)	(0.40)	(0.42)
$\mathbb{1}\{j \neq China\}v_{jg,t-1}$	-0.47***	-0.47***	-0.47***	-0.45***	-0.45***	-0.44***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$\mathbb{1}\{j = China\}v_{jg,t-1}$	-0.36***	-0.36***	-0.38***	-0.36***	-0.37***	-0.32***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
$\mathbb{1}\{j \neq China\}\tau_{jg,t-1}$	-1.73***	-1.73***	-1.75***	-1.93***	-2.23***	-1.56***
	(0.11)	(0.11)	(0.09)	(0.10)	(0.08)	(0.12)
$\mathbb{1}\{j = China\}\tau_{jg,t-1}$	-3.00***	-3.00***	-2.71***	-3.04***	-2.84***	-2.74***
	(0.31)	(0.31)	(0.26)	(0.30)	(0.25)	(0.28)
Shipping Costs$_{jgt}$			-3.20***			
			(0.04)			
Long-run China	-8.23***	-8.23***	-7.21***	-8.36***	-7.71***	-8.67***
	(0.82)	(0.82)	(0.67)	(0.81)	(0.67)	(0.85)
Lon-run Others	-3.70***	-3.70***	-3.71***	-4.27***	-5.00***	-3.55***
	(0.24)	(0.24)	(0.20)	(0.23)	(0.19)	(0.28)
Long-/Short-run China	4.0	4.0	3.79	3.73	3.21	4.45
Long-/Short-run Others	1.52	1.52	1.41	1.72	1.73	1.45
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	733,470	733,470	929,251	808,601	1,024,577	459,202
Adjusted R^2	0.26	0.26	0.29	0.26	0.26	0.25

Notes: The table reports estimates of (1). 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 G.4: Slow adjustment with ECM — robustness, continued

	1974–2008	1974–1988		1989–2008	
	SITC	TSUSA	SITC	HS-8	SITC
$\mathbb{1}\{j \neq China\}\Delta\tau_{jgt}$	-1.82***	-1.89***	-1.99***	-3.47***	-1.52***
	(0.14)	(0.12)	(0.19)	(0.16)	(0.18)
$\mathbb{1}\{j = China\}\Delta\tau_{jgt}$	-2.22***	-1.67***	-2.18***	-2.06***	-1.97**
	(0.40)	(0.28)	(0.42)	(0.56)	(0.88)
$\mathbb{1}\{j \neq China\}v_{jg,t-1}$	-0.48***	-0.78***	-0.66***	-0.61***	-0.55***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$\mathbb{1}\{j = China\}v_{jg,t-1}$	-0.37***	-0.75***	-0.62***	-0.52***	-0.46***
	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
$\mathbb{1}\{j \neq China\}\tau_{jg,t-1}$	-1.49***	-2.26***	-1.78***	-3.00***	-1.45***
	(0.11)	(0.14)	(0.19)	(0.16)	(0.19)
$\mathbb{1}\{j = China\}\tau_{jg,t-1}$	-2.94***	-2.96***	-2.89***	-3.62***	-2.15**
	(0.31)	(0.27)	(0.40)	(0.43)	(0.84)
Shipping Costs_{jgt}	-2.88***	-3.35***	-3.13***	-3.27***	-2.89***
	(0.04)	(0.03)	(0.06)	(0.03)	(0.06)
Long-run China	-7.93***	-3.94***	-4.64***	-7.02***	-4.72***
	(0.81)	(0.36)	(0.62)	(0.83)	(1.83)
Lon-run Others	-3.12***	-2.90***	-2.69***	-4.89***	-2.64***
	(0.24)	(0.18)	(0.29)	(0.26)	(0.34)
Long-/Short-run China	4.36	2.08	2.33	2.02	3.11
Long-/Short-run Others	1.41	1.74	1.23	2.37	1.34
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	733,470	965,641	250,344	1,648,276	478,606
Adjusted R^2	0.29	0.40	0.38	0.32	0.31

Note: This table reports estimates of (1) 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 G.5: NTR-gap elasticity — Pierce and Schott (2016) specification

	NTR Gap		Statutory NNTR		Applied NNTR	
$\mathbb{1}\{^{t>2000}_{j=China}\}NTRGap_g$	-0.67***	-2.07***				
	(0.21)	(0.26)				
$\mathbb{1}\{^{t>2000}_{j=China}\}NNTR$			-0.51***	-1.63***		
			(0.19)	(0.25)		
$\mathbb{1}\{^{t>2000}_{j=China}\}AppNNTRg$					-0.42**	-2.04***
					(0.21)	(0.31)
τ_{jgt}	-3.85***	-3.76***	-3.85***	-3.76***	-3.80***	-3.79***
	(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	703,190	1,245,003	703,804	1,246,136	365,163	740,504
Adjusted R^2	0.84	0.78	0.84	0.78	0.86	0.79

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 G.6: NTR-gap elasticity — TSUSA, 1974–88

	Baseline	All Countries	Avg. NNTR, 1974–79	Tariffs
$\mathbb{1}\{^{t=t'}_{j=China}\}X_g$				
1974	-4.27***	-4.23***	-3.93***	-2.99***
1975	-3.86***	-3.81***	-3.53***	-2.57***
1976	-3.73***	-3.69***	-3.44***	-2.43***
1977	-3.84***	-3.80***	-3.52***	-2.51***
1978	-3.33***	-3.27***	-2.88***	-2.03***
1979	-3.81***	-3.78***	-3.45***	-2.51***
1980	-2.66***	-2.61***	-2.21***	-2.45***
1981	-1.84***	-1.79***	-1.72***	-1.70***
1982	-1.99***	-1.93***	-1.82***	-1.94***
1983	-2.20***	-2.15***	-1.90***	-2.16***
1984	-2.11***	-2.05***	-1.79***	-2.08***
1985	-1.40***	-1.35***	-1.03***	-1.37***
1986	-0.98***	-0.93***	-0.66***	-0.95***
1987	-0.51**	-0.47**	-0.30	-0.50**
τ_{jgt}				-1.89***
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	486,725	500,641	701,141	486,725
Adjusted R^2	0.80	0.80	0.80	0.80

Note: Columns are estimates of (3), 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 (3) 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 G.7: NTR-gap elasticity — HS-8, 1989–2007

	Baseline	No Tariffs	All Countries	Full	Balanced
$\mathbb{1}\{^{t=t'}_{j=China}\}X_g$					
1989	-3.41***	-3.44***	-3.33***	-2.23***	-3.37***
1990	-2.84***	-2.87***	-2.75***	-1.90***	-2.79***
1991	-2.20***	-2.24***	-2.11***	-1.53***	-2.13***
1992	-1.62***	-1.66***	-1.52***	-1.08***	-1.62***
1993	-1.05***	-1.10***	-0.95***	-0.80***	-1.10***
1994	-1.19***	-1.25***	-1.15***	-1.05***	-1.17***
1995	-1.07***	-1.10***	-0.96***	-0.98***	-1.28***
1996	-0.85***	-0.88***	-0.78***	-0.82***	-0.86***
1997	-1.00***	-1.02***	-0.94***	-0.94***	-1.04***
1998	-0.99***	-1.00***	-0.96***	-1.08***	-0.86***
1999	-0.64***	-0.64***	-0.60***	-0.87***	-0.59***
2000	-0.48***	-0.49***	-0.47***	-0.86***	-0.35
2001	-0.51***	-0.53***	-0.49***	-0.97***	-0.34*
2002	-0.29*	-0.31*	-0.28*	-0.58***	-0.22
2003	-0.30*	-0.30*	-0.31**	-0.53***	-0.33*
2004	-0.32**	-0.32**	-0.30**	-0.50***	-0.26
2005	-0.30**	-0.30**	-0.29**	-0.04	-0.16
2006	-0.28**	-0.28**	-0.29**	-0.11	-0.23
2007	-0.16	-0.16	-0.17	0.01	0.02
τ_{jgt}	-3.72***		-4.38***	-4.85***	-4.47***
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	2,029,931	2,029,931	2,322,408	2,939,229	1,210,855
Adjusted R^2	0.77	0.77	0.77	0.77	0.78

Note: All estimates are obtained using (3), 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 (3), 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 G.8: NTR-gap elasticity — alternative samples

	Baseline	All Countries	Full	Balanced	Only U.S.-China
$\mathbb{1}\{^{t=t'}_{j=China}\}X_g$					
1974	-9.72***	-9.64***	-8.44***	-9.57***	-12.26***
1975	-9.48***	-9.48***	-8.01***	-9.31***	-12.52***
1976	-10.63***	-10.58***	-7.75***	-10.48***	-13.01***
1977	-10.47***	-10.42***	-7.54***	-10.32***	-12.45***
1978	-10.57***	-10.54***	-7.33***	-10.41***	-12.90***
1979	-10.82***	-10.80***	-7.22***	-10.68***	-12.65***
1980	-9.19***	-9.18***	-5.87***	-9.03***	-10.94***
1981	-7.32***	-7.31***	-4.47***	-7.23***	-8.99***
1982	-7.63***	-7.61***	-4.57***	-7.38***	-9.25***
1983	-7.66***	-7.62***	-4.19***	-7.35***	-9.31***
1984	-6.33***	-6.32***	-3.68***	-6.56***	-7.67***
1985	-7.28***	-7.25***	-4.40***	-6.84***	-8.40***
1986	-6.95***	-6.97***	-3.99***	-7.00***	-7.77***
1987	-6.27***	-6.23***	-3.66***	-6.25***	-6.84***
1988	-4.95***	-4.90***	-3.10***	-4.71***	-5.53***
1989	-4.13***	-3.99***	-2.49***	-3.65***	-4.97***
1990	-3.41***	-3.24***	-2.42***	-3.04***	-3.90***
1991	-2.69***	-2.59***	-2.04***	-2.46***	-3.05***
1992	-2.19***	-2.09***	-1.68***	-2.45***	-2.44***
1993	-1.77***	-1.62***	-1.26***	-1.95***	-1.78***
1994	-1.56***	-1.60***	-1.50***	-1.97***	-1.71***
1995	-1.58***	-1.53***	-1.72***	-1.67***	-1.65***
1996	-1.48***	-1.45***	-1.67***	-1.30**	-1.54***
1997	-1.57***	-1.62***	-1.65***	-1.36***	-1.69***
1998	-1.26***	-1.28***	-1.65***	-0.52	-1.12**
1999	-1.01**	-1.08**	-1.54***	-0.86*	-1.02**
2000	-0.64*	-0.72*	-1.40***	-0.64	-0.40
2001	-0.15	-0.23	-1.04***	-0.13	0.01
2002	-0.49	-0.56	-1.07***	-0.50	-0.22
2003	-0.89***	-0.99***	-1.30***	-0.61*	-0.54
2004	-0.27	-0.35	-0.85***	-0.22	0.03
2005	-0.52	-0.61*	-0.49*	-0.09	-0.29
2006	-0.26	-0.38	-0.35	-0.48	-0.20
2007	-0.09	-0.13	-0.16	0.21	0.15
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	jt, gj
N	890,190	978,716	1,245,003	544,441	30,163
Adjusted R^2	0.78	0.79	0.78	0.80	0.78

Notes: This table reports the annual NTR-gap elasticities under alternative sample designs. All estimates are obtained using (3). Column 1 corresponds to our baseline estimates reported in Figure 3. 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 G.9: 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	-7.92***	-8.97***	-8.02***	-9.68***	-7.89***	-9.52***
1975	-8.22***	-9.03***	-8.54***	-9.27***	-7.47***	-10.02***
1976	-9.27***	-9.14***	-9.60***	-10.27***	-8.23***	-10.63***
1977	-9.28***	-8.94***	-9.45***	-10.48***	-8.43***	-10.69***
1978	-9.27***	-9.00***	-9.36***	-10.45***	-8.31***	-9.91***
1979	-9.43***	-9.07***	-9.64***	-10.44***	-8.46***	-9.77***
1980	-8.28***	-7.59***	-8.63***	-9.11***	-8.83***	-8.91***
1981	-6.33***	-6.40***	-6.66***	-7.01***	-6.84***	-7.25***
1982	-6.92***	-6.64***	-7.20***	-7.29***	-7.13***	-7.44***
1983	-6.63***	-6.25***	-6.85***	-7.49***	-7.39***	-7.23***
1984	-5.57***	-5.83***	-5.55***	-6.19***	-6.10***	-6.70***
1985	-6.71***	-5.82***	-6.77***	-6.98***	-6.90***	-7.08***
1986	-6.30***	-5.48***	-6.27***	-6.89***	-6.84***	-6.68***
1987	-5.83***	-5.04***	-5.59***	-6.16***	-6.13***	-5.99***
1988	-4.68***	-3.78***	-4.40***	-4.77***	-4.74***	-4.62***
1989	-3.70***	-2.84***	-3.27***	-3.99***	-3.97***	-3.82***
1990	-3.18***	-2.54***	-2.67***	-3.35***	-3.31***	-3.32***
1991	-2.34***	-1.82***	-1.77***	-2.72***	-2.66***	-2.44***
1992	-1.90***	-1.51***	-1.31**	-2.21***	-2.16***	-2.13***
1993	-1.44***	-1.46***	-0.82	-1.78***	-1.72***	-1.66***
1994	-1.22***	-1.40***	-0.62	-1.48***	-1.42***	-1.46***
1995	-1.40***	-0.94*	-0.79	-1.55***	-1.50***	-1.45***
1996	-1.29***	-0.89*	-0.69	-1.47***	-1.38***	-1.38***
1997	-1.43***	-0.98**	-0.83*	-1.48***	-1.46***	-1.37***
1998	-1.31***	-0.55	-0.68	-1.23***	-1.21***	-1.20***
1999	-0.97**	-0.86**	-0.30	-0.95**	-0.94**	-0.85**
2000	-0.65*	-0.53	0.02	-0.65*	-0.65*	-0.27
2001	-0.26	-0.07	0.45	-0.19	-0.19	-0.23
2002	-0.49	-0.09	0.24	-0.47	-0.47	-0.43
2003	-0.88***	-0.33	-0.15	-0.78**	-0.80**	-0.59*
2004	-0.22	0.21	0.51	-0.24	-0.25	-0.37
2005	-0.67**	0.45	0.02	-0.50	-0.50	-0.51
2006	-0.26	-0.16	0.46	-0.26	-0.25	-0.39
2007	-0.13	0.39*	0.61*	-0.14	-0.14	-0.07
log Shipping Cost$_{jgt}$				-3.11***	-3.09***	
τ_{jgt}					-2.56***	
$\Delta\tau_{jg,t+1}$						1.07***
FE	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj	gt, jt, gj
N	891,167	508,882	891,167	890,190	890,190	726,464
Adjusted R^2	0.78	0.80	0.78	0.79	0.79	0.79

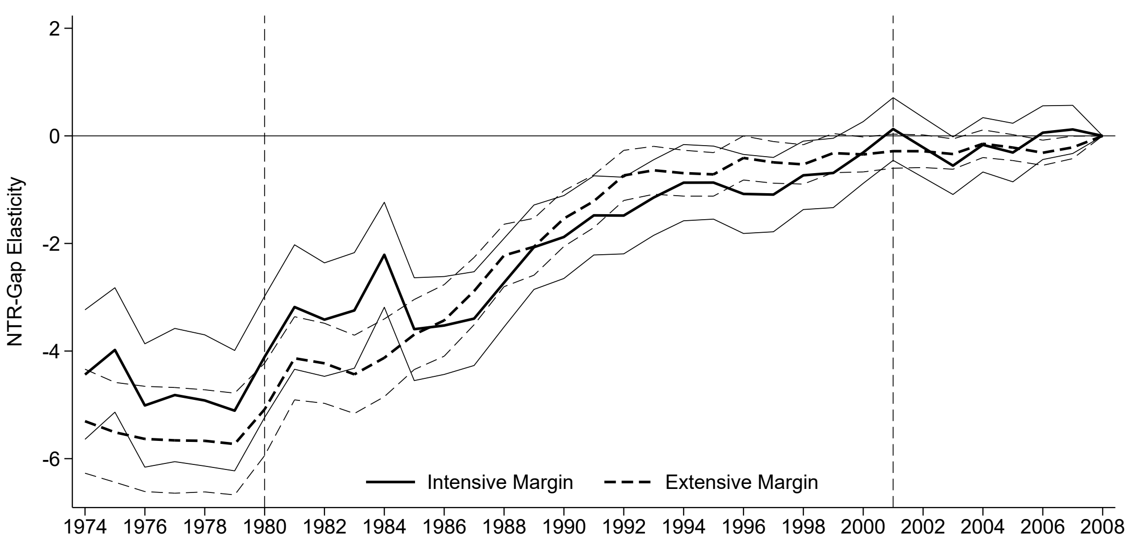
Notes: This table reports the annual NTR-gap elasticities estimated by (3) under alternative specifications of the gap measure, additional trade cost controls, and control for anticipatory effects. Column 1—*NNTR*—defines Gap_g as the NNTR rate. Column 2—*Applied*—defines Gap_g as the applied NNTR rate, 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. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

H Additional figures

Fig. H.1 – Autocorrelation of 1980 tariff changes

Notes: Autocorrelation of 1980 tariff changes, obtained from regressing h-year ($h = [1, 25]$) tariff change relative to 1979 on one-year tariff change between 1980 and 1979. Includes same fixed effects as in (2). 95-percent confidence interval estimated using standard errors clustered at the jg level.

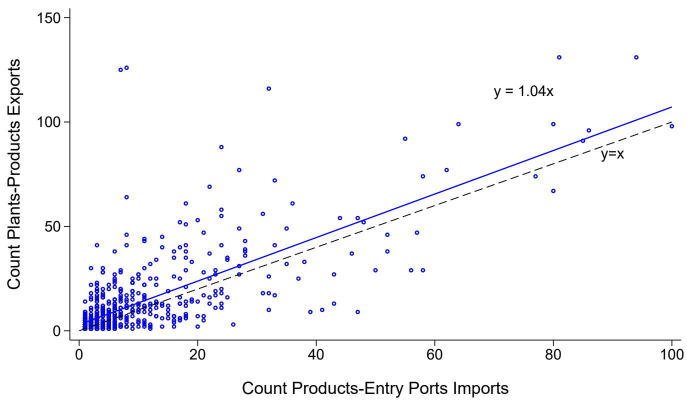
Fig. H.2 – Extensive- and intensive-margin NTR-gap elasticities



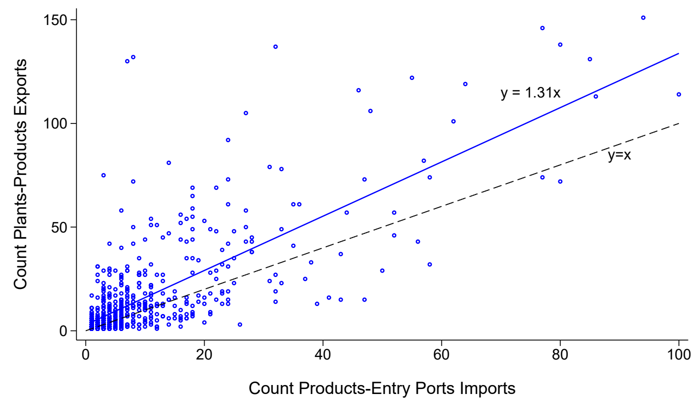
Notes: Estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (3) using measures of extensive (solid line) and intensive margins (dashed line) as dependent variables. 95-percent confidence intervals constructed using standard errors clustered at gj level.

Fig. H.3 – Extensive margin measures in Colombian firm-level export data

(a) Firm count

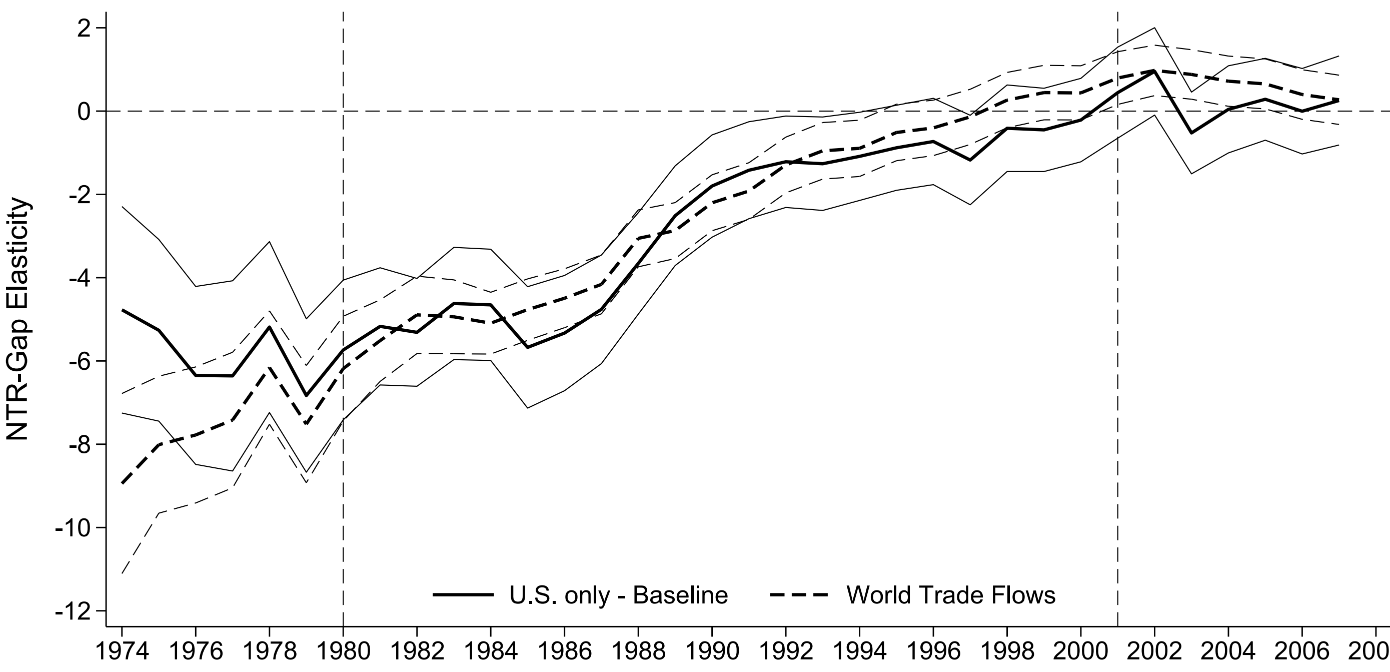


(b) Firm-product count



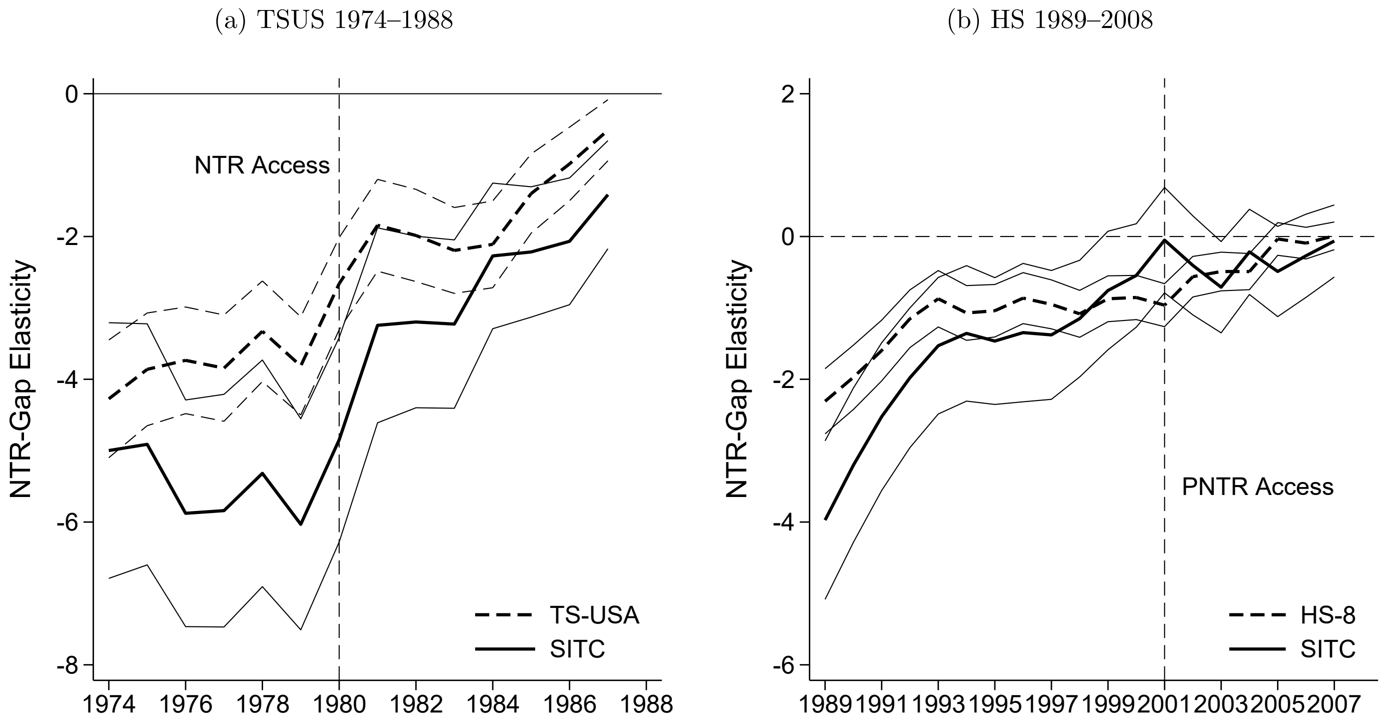
Notes: In both figures the x-axis is the extensive margin measure we use in Figure H.2 applied to U.S. imports from Colombia in 2006: the count of product-district of entry of U.S. imports, where products are defined at the TSUSA 7-digit (1974–1988) and HS 8-digit (1989–2008) level, of each SITC good. 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.

Fig. H.4 – Annual NTR-gap elasticities: China supply factors



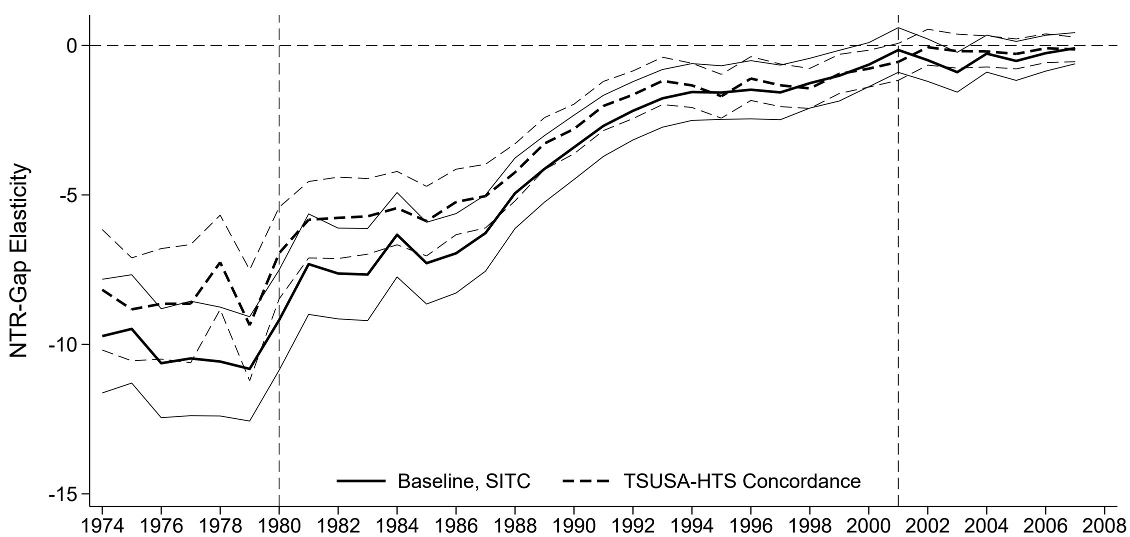
Notes: This figure plots the estimates of $\hat{\beta}_t$ for $t = [1974, 2007]$ from (C.2) using the merged World Trade dataset from Feenstra et al. (2005) (1974–2000) and the BACI Trade Dataset (2000–2008). The standard errors that construct the 95-percent confidence interval are clustered at the jg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

Fig. H.5 – Annual gap elasticities with aggregation at tariff lines



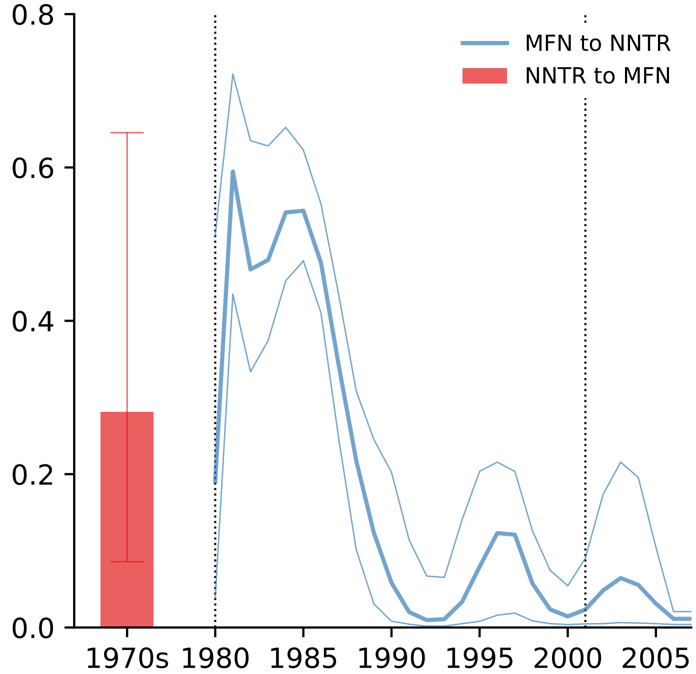
Notes: Left: Estimates of (3) using 7-digit TSUSA classification between 1974 and 1988. Right: estimates of (3) using HS classification between 1989 and 2008. Reference period is last year of each sample period. Estimates from our baseline 5-digit SITC classification are also shown. 95-percent confidence intervals constructed using standard errors clustered at jg level.

Fig. H.6 – Annual NTR-gap elasticities: SITC and TSUSA-HS



Notes: Solid line: baseline NTR-gap elasticity estimates from (3). Dashed line: estimates using a concordance between TSUSA and the HS product schedules from Acosta and Cox (2022). 95-percent confidence intervals constructed using standard errors clustered at jg level.

Fig. H.7 – Trade-policy probabilities: upper and lower bounds

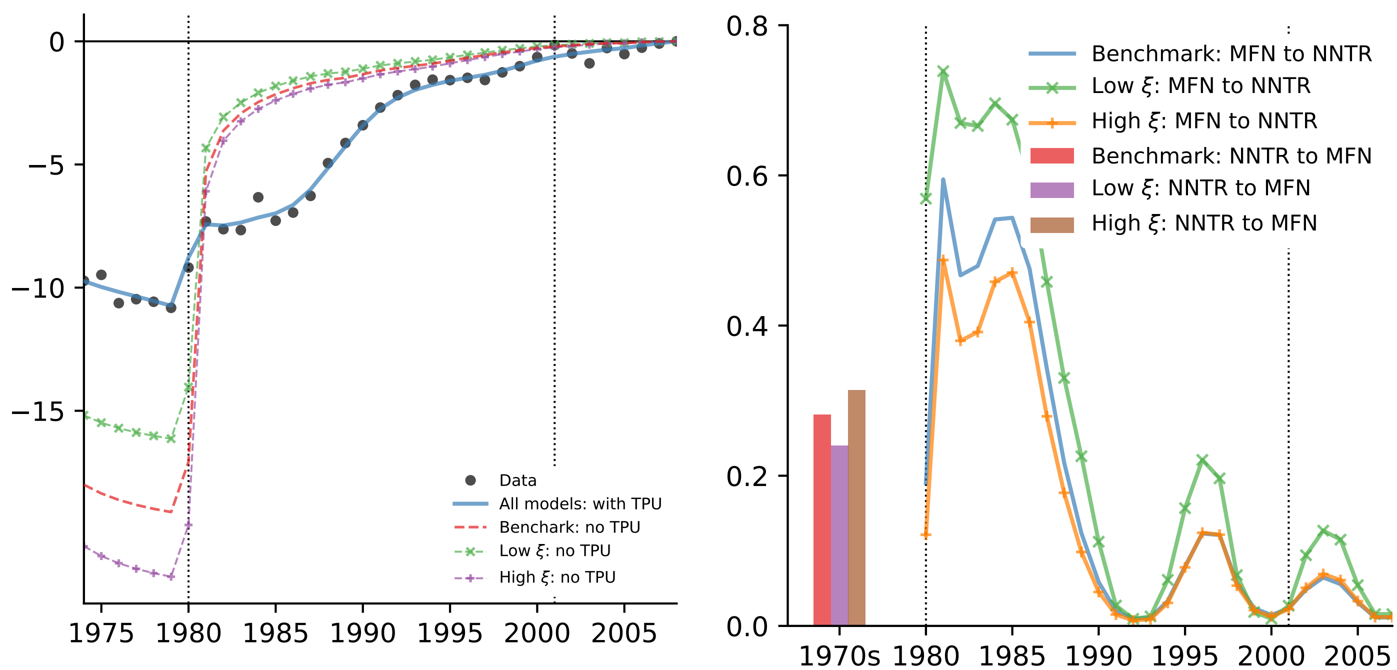


Notes: This figure shows the estimated probabilities of switching policy regimes. Thick lines are the baseline results. Thin lines are estimated by matching the upper and lower confidence intervals shown in Figure 3.

Fig. H.8 – Sensitivity to high iceberg cost (ξ_g)

(a) NTR-gap elasticity

(b) Trade-policy probabilities

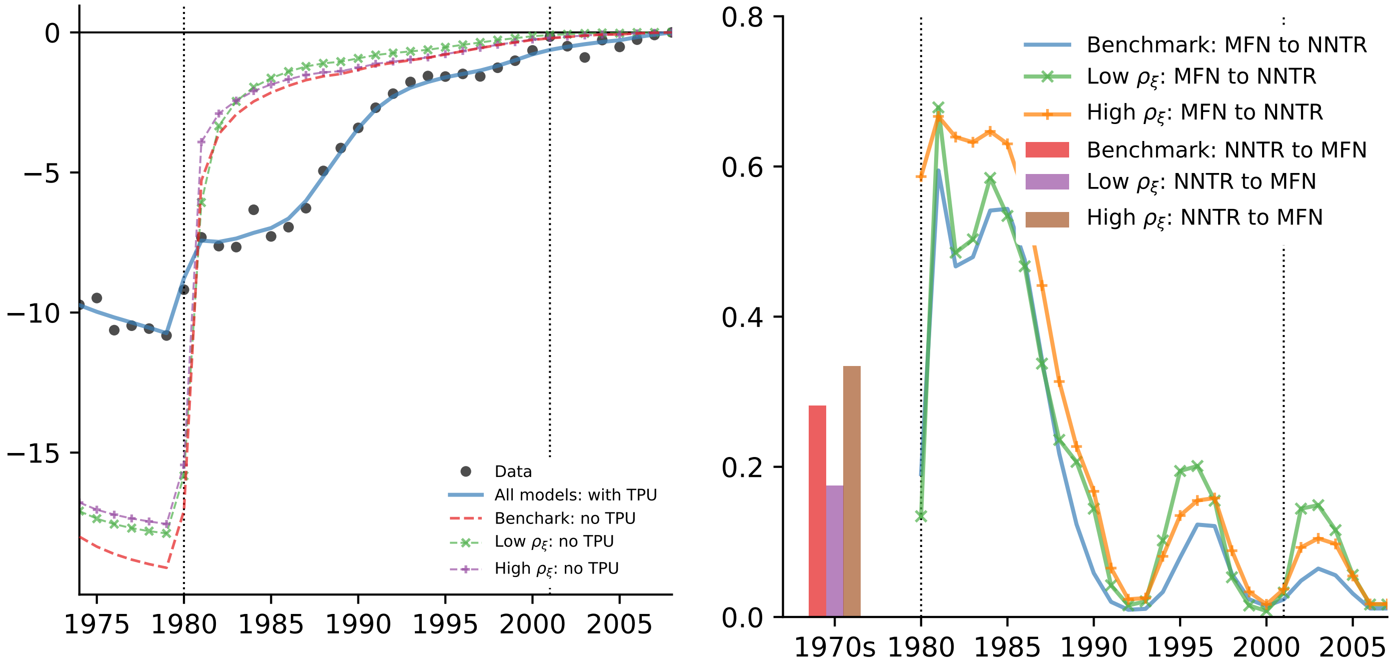


Notes: Panel (a) shows the NTR-gap elasticity in the baseline model vs. alternative calibrations with higher and lower values of ξ_g. Panel (b) shows the estimated trade-policy transition probabilities in these alternative calibrations.

Fig. H.9 – Sensitivity to iceberg cost transition probability (ρ_ξ)

(a) NTR-gap elasticity

(b) Trade-policy probabilities

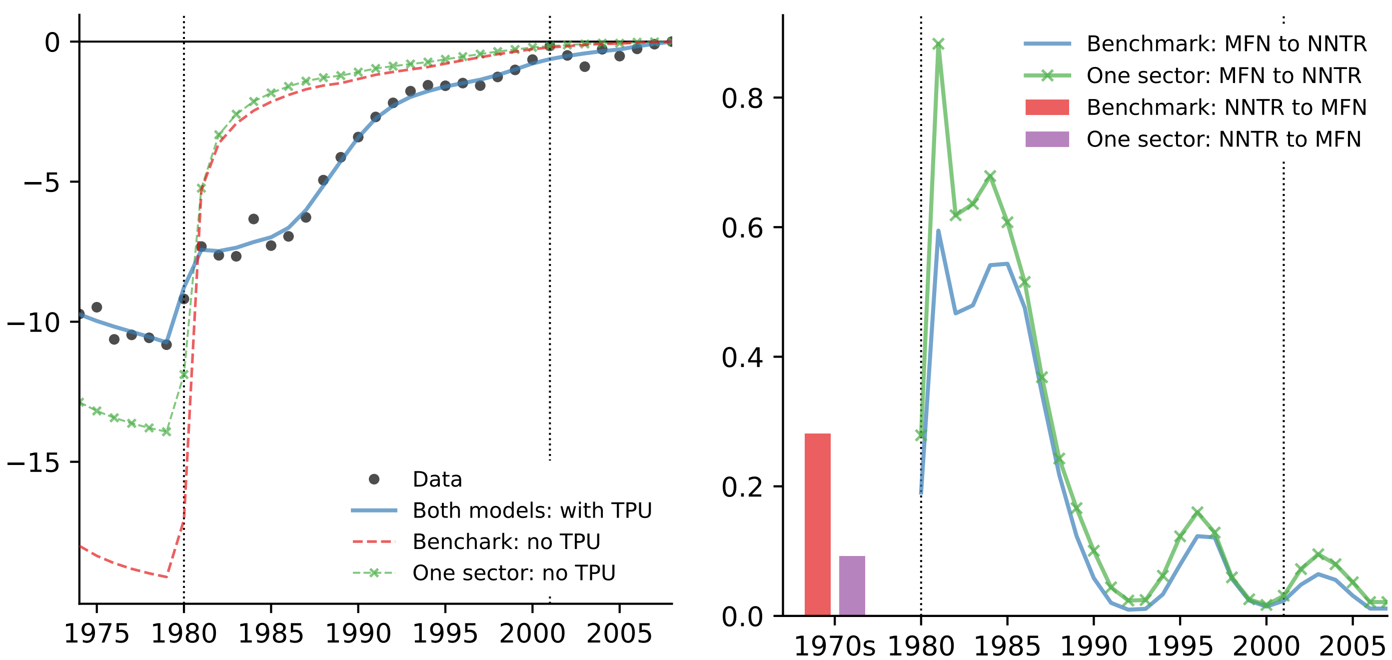


Notes: Panel (a) shows the NTR-gap elasticity in the baseline model vs. alternative calibrations with higher and lower values of ρ_ξ. Panel (b) shows the estimated trade-policy transition probabilities in these alternative calibrations.

Fig. H.10 – Benchmark model vs. one-sector model

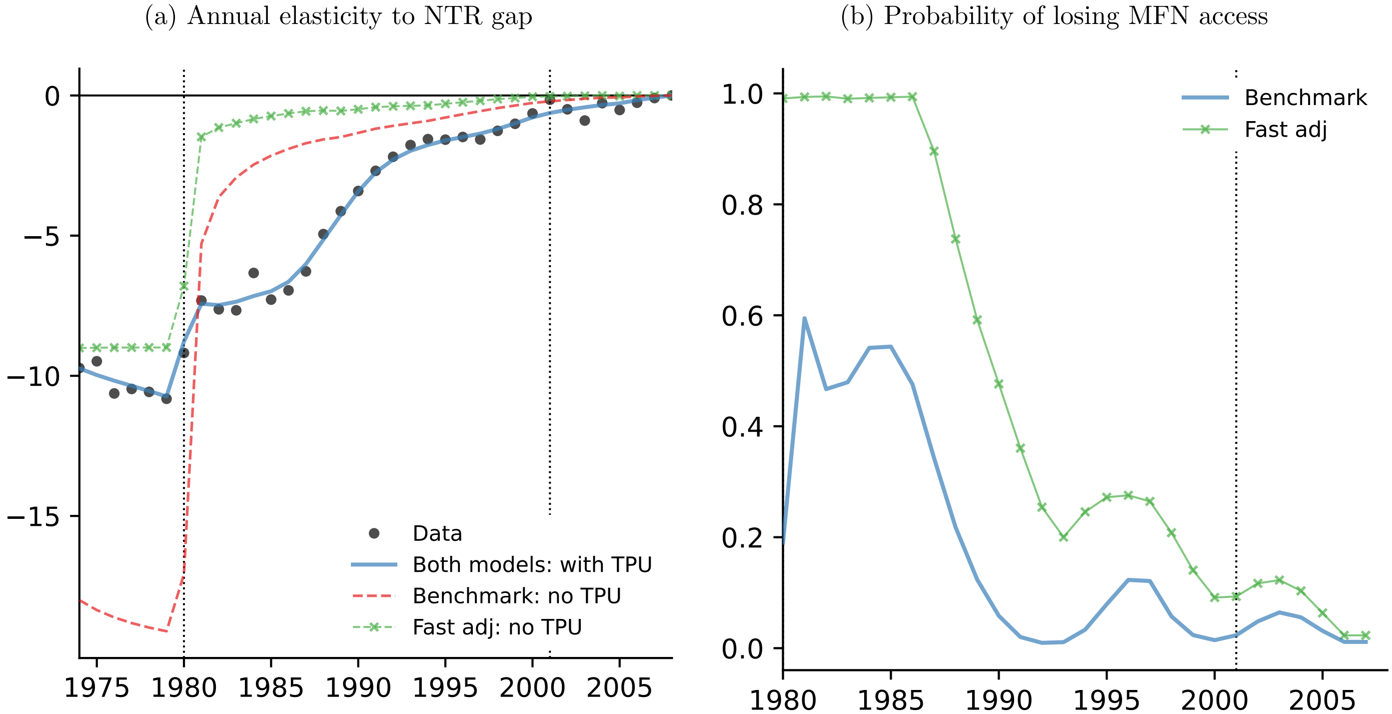
(a) Annual elasticity to NTR gap

(b) Trade-policy probabilities



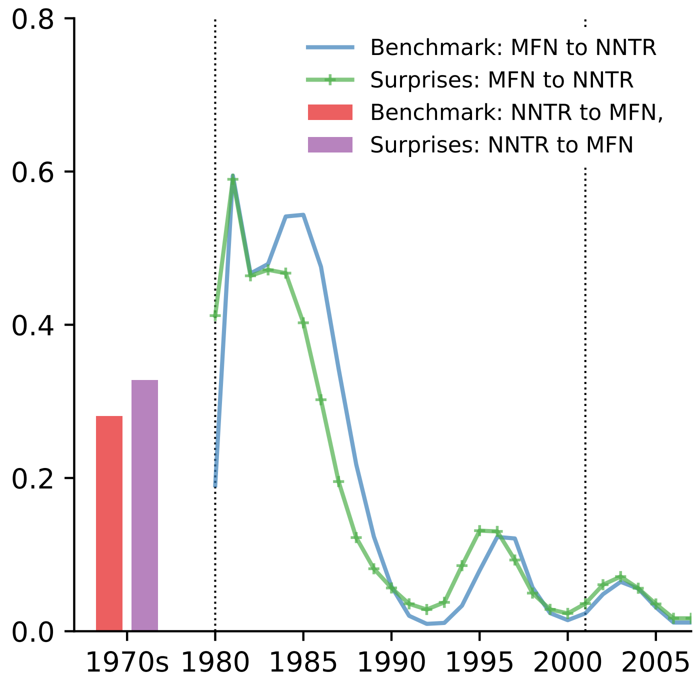
Notes: Results for benchmark vs. one-sector model. Panel (a): NTR-gap elasticities. Panel (b): trade-policy transition probabilities. In one-sector model, all goods have same demand elasticity, non-tariff trade costs, and productivity dispersion; parameters calibrated to match aggregate exporter-dynamics statistics instead of sector-level statistics.

Fig. H.11 – Benchmark model vs. fast-adjustment model



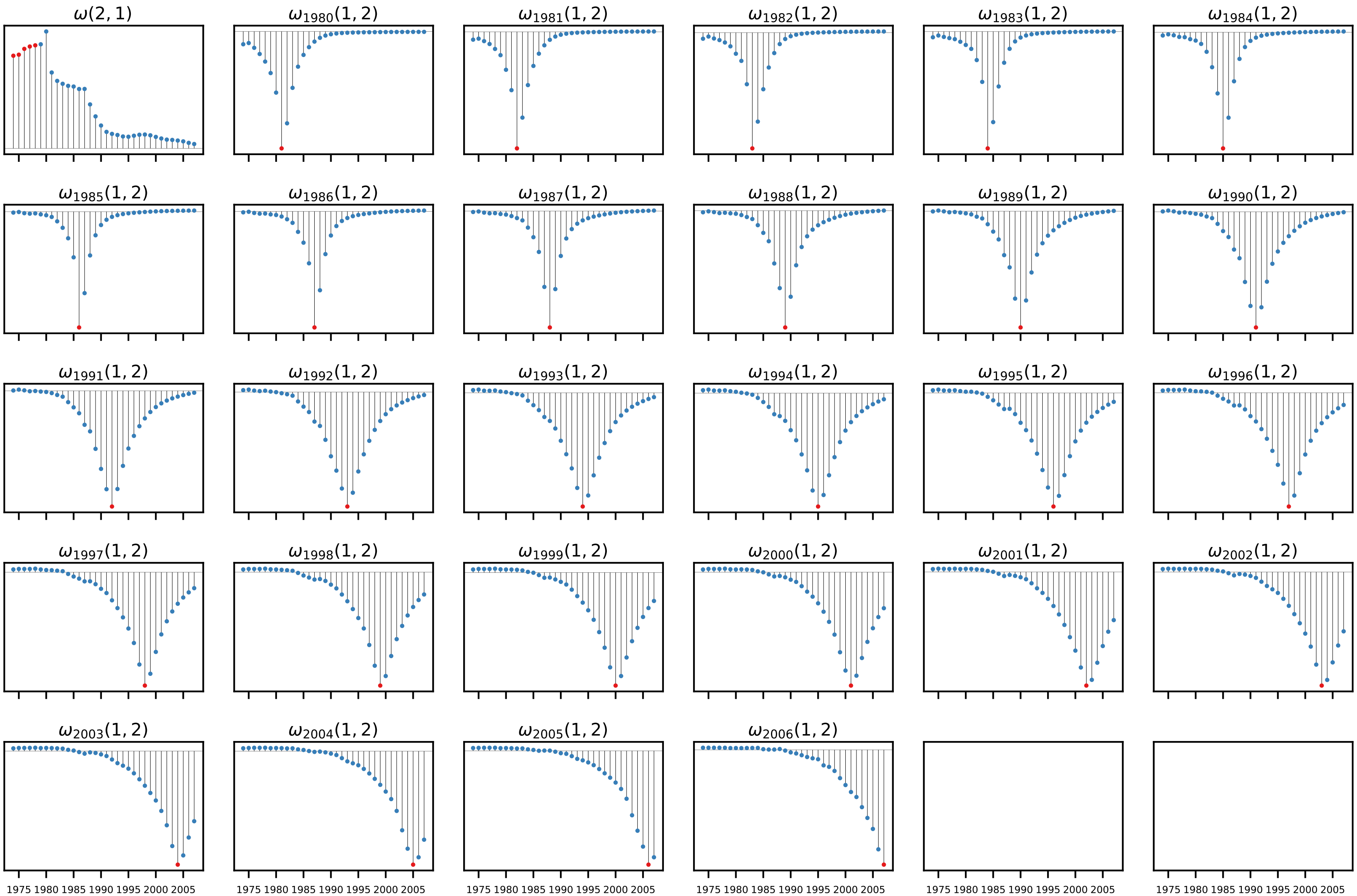
Notes: Results for benchmark vs. fast-adjustment model. Panel (a): NTR-gap elasticities. Panel (b): trade-policy transition probabilities. In the fast-adjustment model, there are no idiosyncratic shocks to firms' productivities or variable trade costs.

Fig. H.12 – Trade-policy probabilities: unanticipated changes



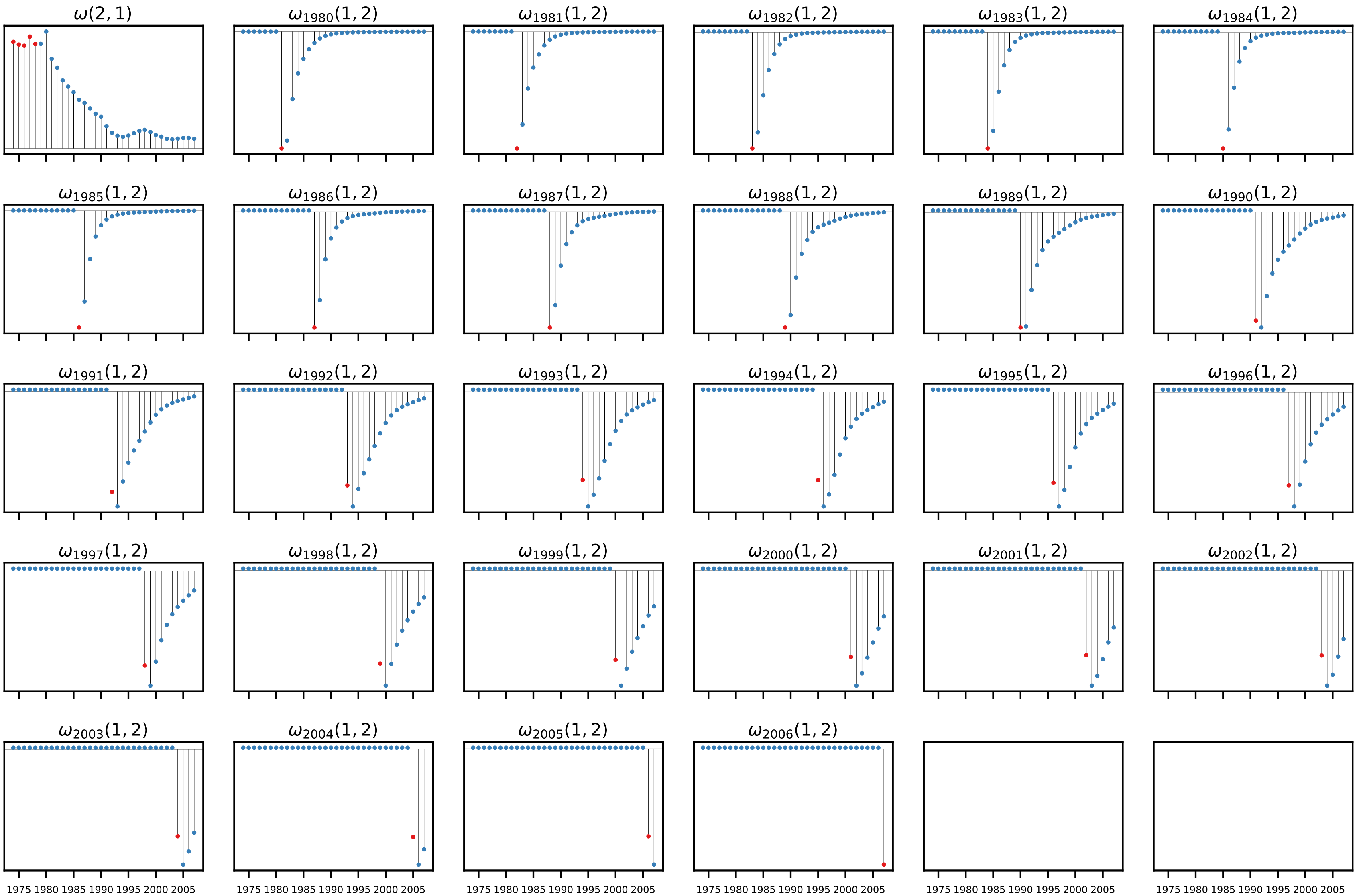
Notes: Estimated probabilities of switching policy regimes in benchmark vs. surprises model.

Fig. H.13 – Sensitivity of NTR-gap elasticities to probabilities:
benchmark model



Notes: Each panel shows how NTR-gap elasticities in benchmark model change when single trade-policy transition probability is increased by 1 percentage point. First panel: probability of gaining NTR access, $\omega(0,1)$. Remaining panels: probability of losing NTR access, $\omega_t(1,0)$. Red dots: elasticities used for identification. Blue dots: elasticities not used for identification.

Fig. H.14 – Sensitivity of NTR-gap elasticities to probabilities: surprises model

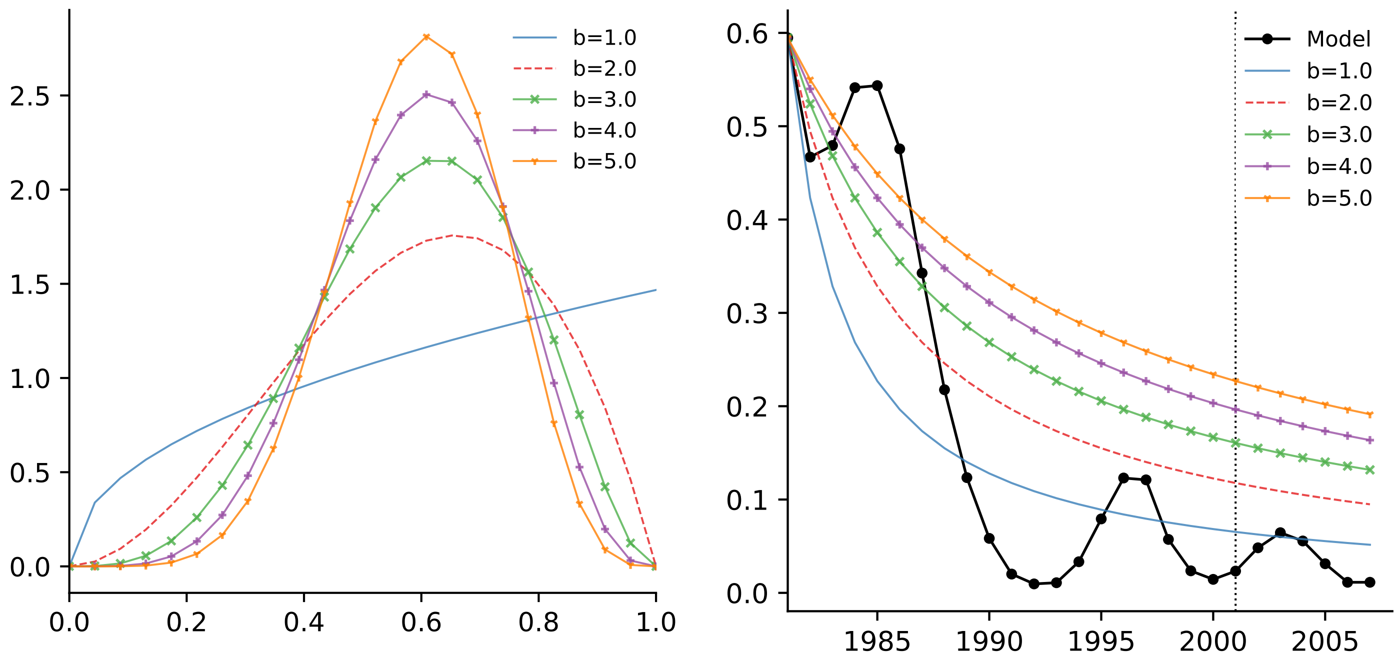


Notes: Each panel shows how NTR-gap elasticities in benchmark model change when single trade-policy transition probability is increased by 1 percentage point. First panel: probability of gaining NTR access, $\omega(0,1)$. Remaining panels: probability of losing NTR access, $\omega_t(1,0)$. Red dots: elasticities used for identification. Blue dots: elasticities not used for identification.

Fig. H.15 – Model-implied expectations vs. Bayesian learning

(a) Prior PDF

(b) Posterior mean

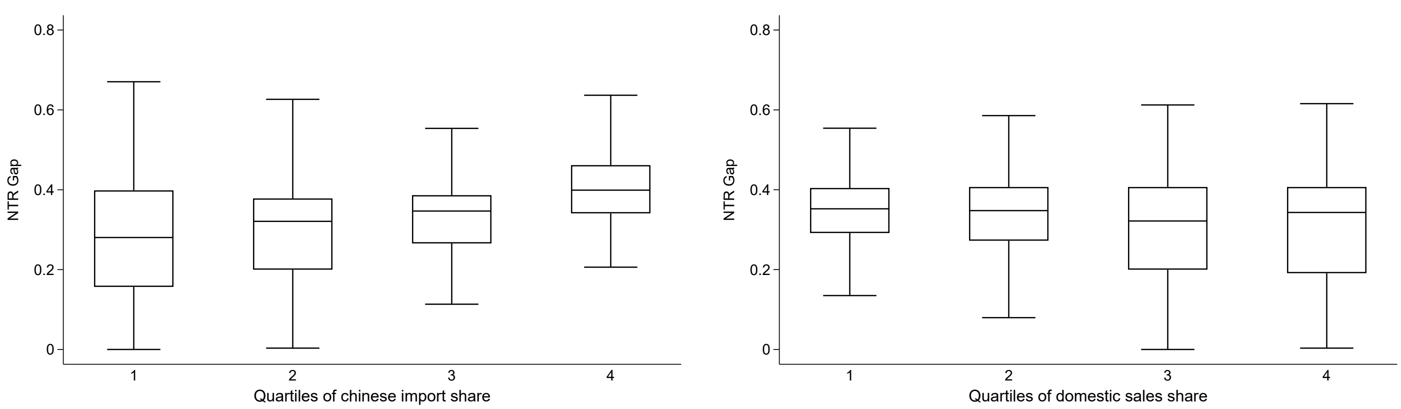


Notes: Panel (a) shows Bayesian prior-belief distributions about the probability of losing NTR status. All priors have the same mean as the model-implied estimate for 1980. Panel (b) shows mean posteriors from 1980–2008 against the model-estimated probabilities.

Fig. H.16 – Variation in the NTR gap across industries versus variation in Chinese import shares

(a) Chinese import share quartiles

(b) Domestic sales share quartiles

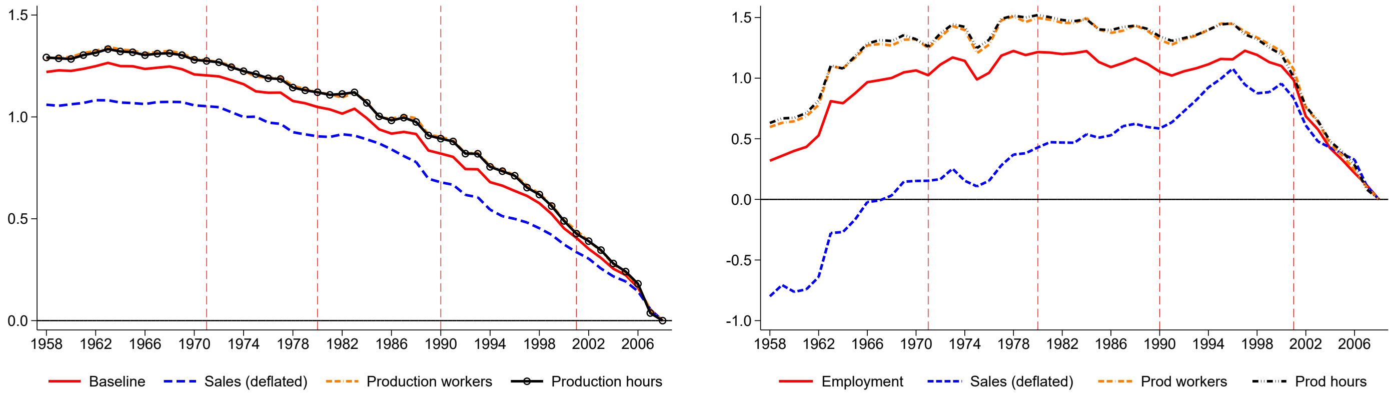


Notes: This figure plots the distribution of NTR gaps across different quartiles of China's import share of industry absorption (Panel a) and the domestic market share of U.S. firms in total industry sales (Panel b). Both are simple averages during 1995—1999.

Fig. H.17 – Alternative measures of substitution

(a) Baseline estimate

(b) Reduced form elasticities

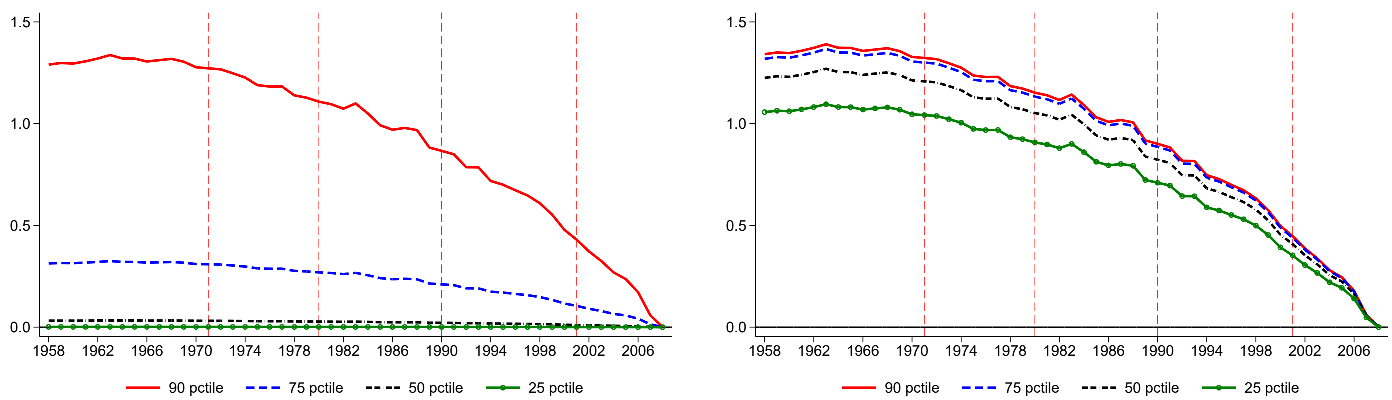


Notes: This figure plots versions of Figure 7 with sectoral deflated sales, production workers, and sectoral production hours as the dependent variables instead of total sectoral employment.

Fig. H.18 – Sectoral employment effects by industry import and export share

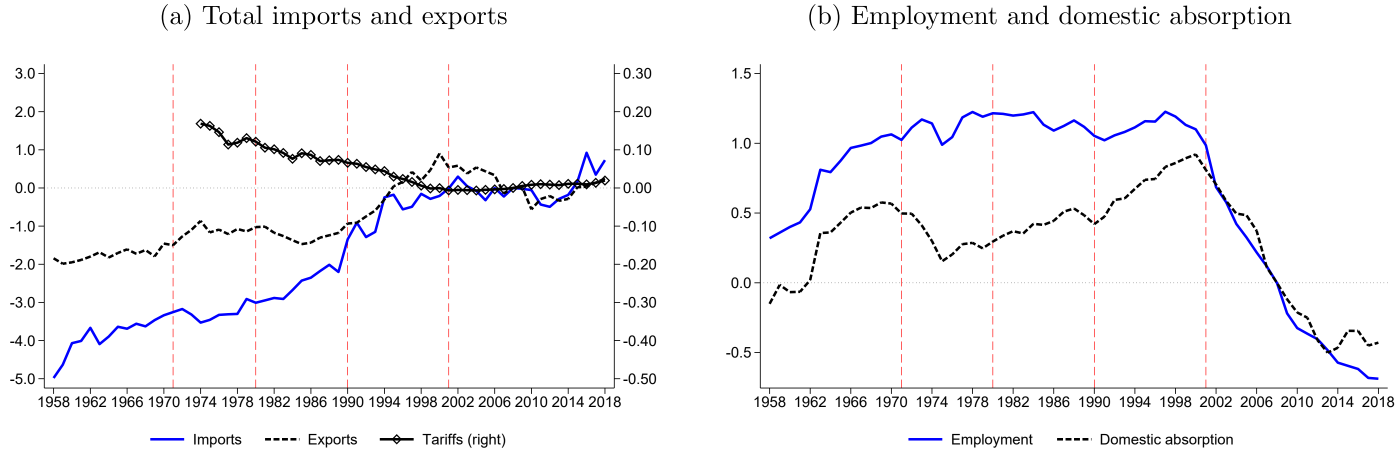
(a) Chinese import share percentiles

(b) Domestic sales share percentiles



Notes: This figure highlights the industry-level heterogeneity in the annual coefficients from (20). 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.

Fig. H.19 – Elasticity to the NTR gap



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.

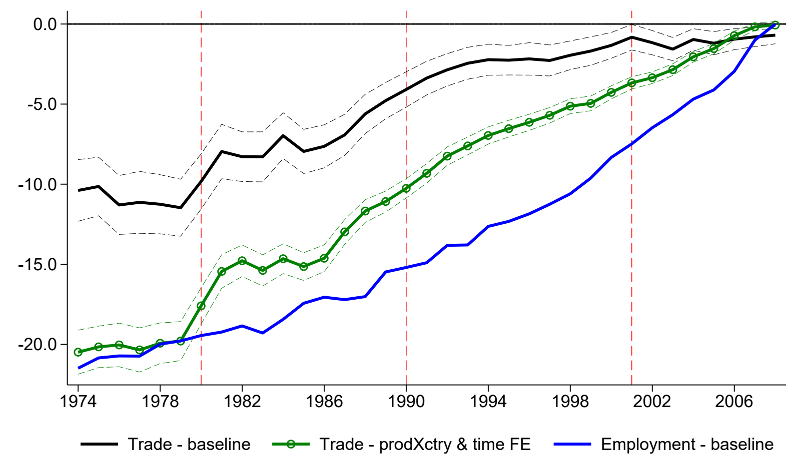


Fig. H.20 – NTR-gap elasticities estimated using trade vs. employment data. Solid line: using employment data. Dashed line: using trade data with product-time (δ_{gt}), product-country (δ_{jg}), and country-time (δ_{jt}) fixed effects. Dotted line: using trade data with only product-country (δ_{jg}) and time (δ_t) fixed effects. Shaded areas: 95-percent confidence intervals.