Recovering Credible Trade Elasticities from Incredible Trade Reforms*

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January 2025

Abstract

We study how trade-policy dynamics affect the dynamics of trade volumes and the implications of these effects for estimates of the trade elasticity. We use data on U.S. imports and trade policy from 1974-2017 for China and Vietnam—the countries with the largest import growth and the largest tariff reductions over the last fifty years—and a heterogeneous-firm dynamic trade model to recover the dynamic path of the trade elasticity following an unanticipated, permanent tariff change. We estimate a short-run trade elasticity of about three and a long-run trade elasticity of about 15, and find that it takes about three to five years to close half of the gap between the current and long-run levels of trade. We argue that the expected dynamics of future trade policy before and after these reforms biases down reduced-form estimates of the long-run trade elasticity and biases up estimates of the short-run elasticity. We argue that these measurement issues are even more problematic for other trade reforms, especially those within the Normal Trade Relations (NTR) regime that constitute the majority of the data.

JEL Classifications: F12, F13, F14

Keywords: Trade elasticity, trade dynamics, trade policy uncertainty

*We thank seminar audiences at the University of Rochester, St. Louis Federal Reserve, Syracuse University, 2024 ASSA Meetings, and the 2024 Canadian International Trade Study Group for comments. Kazuma Inagaki provided excellent research assistance. Alessandria and Ruhl are grateful for support from the National Science Foundation (award #2214852).

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1 Introduction

The elasticity of trade volumes to trade policy is one of the most important concepts in international economics. It determines how trade reforms affect substitution between domestic and foreign goods, reallocation of workers across sectors, and aggregate welfare. Despite its importance, there is no clear consensus on the magnitude of the trade elasticity for two key reasons. First, trade responds gradually to policy changes, so the elasticity depends on the time horizon over which it is measured. Second, the response depends on expectations about future policy, which means that the same observed policy change can generate many different outcomes.

Ideally, one would like to estimate a time-series for the trade elasticity following an unanticipated, once-and-for-all trade liberalization, but we argue that such reforms do not exist in the data. We recover this path by using data from two unique trade reforms—the United States granting Normal Trade Relations (NTR) to Vietnam and China—as inputs to a structural model with a time-varying trade policy process. We find that the short-run trade elasticity is about three, while the long-run trade elasticity is about 15. One-half of the gap between the short and long-run elasticity closes in 5 years.

Building on the insights of Baldwin (1986) and Baldwin and Krugman (1989) that trade is driven by firm-level investments in market access, a large literature in structural microeconomics finds the structure of trade costs makes export participation a forward-looking decision (see, for example, Roberts and Tybout, 1997; Das et al., 2007). A common finding in this work is that entering and growing in export markets involves repeated up-front investments with back-loaded returns, which implies that trade responds gradually to policy changes (Ruhl and Willis, 2017; Alessandria et al., 2021). It also implies that trade depends on the underlying policy process—and changes to that process—as well as the observed sequence of realizations. For example, Ruhl (2008) argues that trade responds less to temporary shocks than permanent shocks, and Handley and Limao (2014) argue that uncertainty about future policy depresses trade in the present. Thus, it is difficult to measure the trade elasticity using data alone, as one must separately identify the lagged effects of past reforms and the effects of expectations about future reforms. Our approach, which builds

on Alessandria et al. (2025), is to use the data to estimate a structural model by indirect inference, and then measure the trade elasticity by simulating the desired unanticipated, once-and-for-all liberalization.

In the first part of the paper, we use data on U.S. imports from 1974–2018 to estimate reduced-form trade responses to tariff changes. Throughout this period, the United States has maintained two main tariff schedules: Non-Normal Trade Relations (NNTR) and Normal Trade Relations (NTR). The NNTR schedule was set by the 1930 Smoot-Hawley Tariff Act and is considered exogenous to any specific trading partner (Handley and Limão, 2017; Pierce and Schott, 2016). The NTR schedule has evolved gradually through multilateral negotiations, making it largely exogenous as well. Additionally, some countries have had free trade with the United States under Preferential Trade Agreements (PTAs) or the Unilateral Trade Preference Program (UTPP). This structure implies that there are two kinds of tariff changes: (i) those that occur when countries shift from one schedule to another, leading to simultaneous tariff changes across many goods at once; (ii) and those that occur when the schedules themselves change, leading to good-specific changes in tariffs that are common across the countries that trade under that schedule. We compare the former with the latter, both in terms of the dynamics of tariffs themselves but also in terms of the effects on trade. Schedule transitions are rare, but cause large, highly-persistent changes in tariffs and have significantly higher long-run trade elasticities than typical estimates in the literature. Within-schedule tariff changes are common but small and transitory, and have low long-run trade elasticities that are similar to typical estimates. We also find that some schedule switches, namely entering/exiting PTAs or the UTPP program, have higher short-run trade elasticities than other tariff changes, which is likely because these switches are more highly anticipated.[1].

We also compare two case studies of transitions between the NNTR and NTR regimes: China and Vietnam. The case of China has been studied extensively. China gained conditional MFN status in 1980, but there was considerable uncertainty about whether this status would be revoked until its 2001 accession to the World Trade Organization (WTO)

[1]PTAs are the product of years of public negotiations and involve tariff changes that are often phased-in gradually. Countries typically leave the UTPP program when they cross predetermined income per capita thresholds, which is also predictable.

(Pierce and Schott, 2016; Handley and Limão, 2017; Alessandria et al., 2025). Vietnam went through a qualitatively similar experience: it gained conditional MFN status in 2002 and permanent status when it joined the WTO in 2007. We document several important differences in the growth of U.S. imports from these two countries that suggest different degrees of anticipation and uncertainty. Both countries have similar long-run trade elasticities that are substantially larger than the average across all regime changes (and far larger than the within-regime elasticity). There are also material differences in trade dynamics between the two countries. Namely, Vietnam's short-run elasticity is higher and it converged to the long-run faster. This suggests that Vietnam's MFN access was more anticipated and was initially viewed as more credible. We argue that this is consistent with the historical evidence. China was one of the first ex-communist countries to gain MFN status and its relationship with the United States was fraught with political tensions. Vietnam, on the other hand, was one of the last countries to gain this status, and, by the time it did so, only one other country had ever lost it.

In the second part of the paper, we use the approach developed in Alessandria et al. (2025) to estimate time-varying regime transition probabilities in a structural model to match the different paths of trade dynamics for China and Vietnam. The model has two key ingredients: sunk exporting costs and export capacity that grows stochastically with a firm's tenure as an exporter. The first ingredient leads firms to make forward-looking export participation decisions that depend on the entire future path of trade policy. The second ingredient allows the model to accurately capture the rate at which firms discount future profits and leads to gradual adjustments in aggregate trade. For each country, we calibrate the model to match two sets of empirical moments about trade dynamics: microeconomic facts from firm-level panel data; and the path of the NNTR-gap elasticity estimated in the first part of the paper. The identification of the policy regime transition probabilities comes from the latter. When the country trades under the NNTR schedule, an increase in the probability of gaining NTR status has a larger positive effect on exports of goods with high NNTR gaps than low-gap goods, which raises the gap elasticity. When the country trades under the NTR schedule, an increase in the probability of switching back to the NNTR regime has a larger negative effect on high-gap goods, which lowers the gap elasticity.

Mirroring the results from our reduced-form analysis of these case studies, we find similarities in the beginning and end of sample, but some noticeable differences around the transition to NTR. In both cases, we find a low initial probability of gaining NTR access, between five and 15 percent. For Vietnam, the likelihood of switching to NTR regime for gradually rises to almost 30 percent several years before this switch actually occurs, whereas the probability drifts down slightly for China. After gaining NTR access, China's probability of reverting back to NNTR temporarily increases before beginning to fall, whereas in Vietnam it begins to fall right away. At the end of the observation period, the probability of losing NTR access is low but nonzero, even after WTO accession. These results indicate that the reduced-form trade-elasticity estimates are indeed biased by anticipation, both in the short run (especially for Vietnam) and in the long run.

To evaluate the magnitude of these biases, we perform the canonical unanticipated, once-and-for-all reform in our model. We solve the model under the counterfactual assumption that each country was initially in a steady state with NNTR tariffs and no chance of moving to the NTR regime, and then NTR is granted unexpectedly with no chance of reversal. We recover very similar trade-elasticity dynamics for both countries: a long-run trade elasticity of about 15, a short-run elasticity of 3, and a speed of adjustment such that the gap between the current and long-run levels of trade is closed every three to five years. This indicates that our long-run reduced form estimates for China and Vietnam are biased downward by about 25 percent, and that our short-run estimate for Vietnam is biased upwards by about 100 percent.

Our paper relates to several papers that estimate the dynamics of the trade substitution around changes in trade policy. A key challenge in much of this work is the most-favored-nation principle, which implies that there is a dearth of variation in tariffs across countries that can be used to identify these dynamics. Thus, several papers have leveraged the tariff variations from PTAs to estimate trade dynamics. Baier and Bergstrand (2007) use a panel of PTAs to identify the dynamics of trade growth that start prior to the signing of the PTA and stretch 10-15 years following the PTA. Baier et al., 2014; Kehoe and Ruhl, 2013; Alessandria and Avila, 2023 show that trade agreements lead to a large increase in the extensive margin of trade measured by the range of products or firms exporting. Focusing more narrowly on

NAFTA, Khan and Khederlarian, 2021 estimate a short-run elasticity of 2.7 and a 7 year response that is close to 9 for trade with Mexico. Similarly, Besedes et al., 2020 find delayed effects of import growth in Mexico and Canada for goods that have immediate tariff cuts or changes in trade policy risk.

A potential concern with this approach is that the path of tariffs (either within-schedule changes or movements of countries across schedules themselves) may be endogenous. To account for this endogeneity, Boehm et al. (2023) focus on the effects of within-NTR tariff changes on minor trading partners, which are unlikely to play a material role in determining countries' strategies in setting NTR tariff schedules. Using COMTRADE data on trade and applied tariffs, they estimate trade elasticities that are much lower than our estimates: less than one in the short run and less than two in the long run.[2] However, when we focus on these within-NTR tariff changes, we recover similarly low estimates. This highlights an important external validity concern: estimates of trade elasticities for small, transitory tariff changes within the NTR schedule are likely to be poor predictors of the effects of unanticipated, large, and persistent reforms like PTA formation and NTR access grants.

More broadly, several papers estimate the dynamic response to changes in the exchange rate or relative prices at the aggregate level or more aggregated industry-level Hooper et al., 2000; Gallaway et al., 2003. Much of this work too finds a delayed response, although estimates of the trade elasticity tend to be much lower. For instance, Alessandria and Choi (2021) use U.S. net trade flows, relative prices and expenditures to estimate a quarterly short-run elasticity of 0.2 and a long-run elasticity of close to 1.1, although they also show that purely empirical approaches that do not control for the source of the shock can substantially bias these estimates downward. Our paper also relates to a growing literature on the effects of future trade policy on the current level of trade reviewed by Handley and Limão (2022). This literature seeks to identify trade growth from changes in the expected persistence of a trade policy, but largely abstracts from measuring how these changes biases estimates of the trade elasticity.

[2]Recently, Teti (2023) demonstrates that the tariff data in COMTRADE is plagued by measurement errors that severely biases these estimates downward.

2 Empirical evidence on trade-policy dynamics

We begin by documenting several facts about the dynamics of trade policy. We focus on the United States owing to the availability of high-quality data over a long period. We emphasize that unanticipated once-and-for-all (UOAFA) trade reforms—the canonical reform studied in most models—are extremely rare. Most tariff changes are highly transitory, and more persistent reforms are often characterized by anticipation due to phaseouts or long negotiation periods. We provide an overview of the structure of the U.S. tariff schedule along with a brief history, analyze the properties of tariff changes, and argue that the grants of NTR status to China and Vietnam are the closest empirical matches to UOAFA reforms.

2.1 Overview and History

The United States has a unique tariff structure that introduces some important tariff variation that facilitates estimating trade responses. Specifically, imports into the U.S. generally enter under one of four tariff regimes:

1. Normal Trade Relations (NTR, or Column 1)
2. Non-Normal Trade Relations (NNTR, or Column 2)
3. Unilateral Trade Preference Program (UTPP)
4. Preferential Trade Agreements (PTA)

For most of its trading partners, the United States uses two tariff schedules, NNTR and NTR. The NNTR schedule is largely determined by the Smoot-Hawley Act of 1930 and is often treated as a source of exogenous variation in product-level tariffs (Pierce and Schott, 2016; Handley and Limão, 2017). The NTR schedule, formerly the Most Favored Nation schedule, has been adjusted over time through multilateral GATT rounds involving successively larger groups of developed countries. Three major GATT rounds—Kennedy, Tokyo, and Uruguay—had large impacts on tariffs, involved long negotiations, and were implemented in phaseout periods of 5–8 years. Prior to the Trade Agreements Extension Act of 1951, trade policy followed the most-favored-nation principle and all U.S. trading partners, save those for which the U.S. did not trade,[3] received the NTR rate. With the 1951 Trade Act, all

[3]Under the Trading with the Enemy Act of 1917, the United States imposed a trade embargo on China and North Korea in December of 1950.

non-market economies, except Yugoslavia, were transitioned from the NTR to the NNTR schedule. Of the 29 countries, which includes successor countries, subject to NNTR under the 1951 trade act or subsequent legislation, only North Korea and Cuba have NNTR status at the end of our sample but trade with both countries are subject to an embargo.[4]

With the Trade Act of 1974, the UTPP regime came into existence with the General System of Preferences (GSP). This program provided tariff-free access to the United States for some products from less-developed countries. The program has lapsed several times, and countries may be graduated from the program upon reaching thresholds in level of development or market share in a product, their accession to other trade agreements, or other circumstances.[5] While GSP is the best-known temporary trade program, there are several other programs, which we describe in Table A1. Another key provision of the 1974 Trade Act was the Jackson-Vanik Amendment, which created a path for non-market economies to gain access to NTR rates subject to Congressional approval and annual renewal. Finally, PTAs came into being in 1985, with the largest U.S. FTA being the Canada-U.S. FTA in 1989 followed by North American FTA in 1995. As of 2017, the United States has 20 FTAs, as listed in Table A3. PTAs often involve domestic content requirements leading some products to be imported under both PTA and NTR rates during the same year.

2.2 Data

Our analysis is based on annual customs data from the Census from 1974 to 2017. We start in 1974 as this is the first year that applied tariffs are available digitally from the National Archives. We continue until 2017 to avoid effects from the U.S.-China Trade War. The data for 1974 to 1988 are from the National Archives and Records Administration, while the data for the remaining years are from Schott (2021). To study trade over such a long horizon, we aggregate the data to the SITC 5-digit level.[6] For these SITC goods we use applied tariffs

[4]For a summary of U.S. trade policy regimes see CRS Report for Congress (2005).

[5]For instance, former non-market economies in the European Union lost access to GSP rates when they joined the European Union.

[6]A challenge to analyze trade over a long time period is that the classification of products into tariff lines changed in 1989, when the United States switched from the 7-digit TS-USA classification to the 8-digit Harmonized System (HS). Our approach is to use data aggregated to the 5-digit Standard International Trade Classification (SITC) level, for which concordances with the TS-USA and HS schedules are both readily available. The concordance between the TS-USA schedule and the SITC schedule is from Feenstra (1996) and the concordances with the HS schedule and its successive revisions are from United Nations Trade

rather than statutory tariffs. To account for changes in country definitions, particularly with the fall of the Soviet Union and its satellites, we merge countries that at some point in our sample period have been united, e.g. the Czech Republic and Slovakia after 1989, and East and West Germany before 1990. Our final sample is an unbalanced panel with 44 years, 163 countries, 2,032 goods, and 2,279,579 observations.

Our unit of observation is the triplet jgt, where j denotes the exporting country, g denotes an SITC 5-digit good, and t denotes the year. Our main variables of interest are the log FOB import value, denoted by v_{jgt}, and the applied tariff rate, denoted by τ_{jgt}. We compute the latter as the ratio of total duty charges to the FOB import value. To classify triplets into different statutory tariff regimes, we use information on the rate of provision and country sub-code. The rate of provision and the country sub-code indicate the type of duty and special trade agreement under which a good from a country is imported, respectively.[7] As our data combines products that come in under different rates of provisions, we assign products to rates of provision based on the most common provision by value in a SITC-good.

Figure 1 plots some features of trade and tariffs by tariff regime. Panel (a) presents the share of U.S. non-oil imports, by current regime of the importer, relative to the share in the NTR regime. Panel (b) shows the path of tariffs for those same countries.[8] In 1974, nearly all U.S. imports were under NTR. In 1976, with the introduction of the GSP program, there is a discrete jump in the share of countries that gain UTPP and its associated zero tariffs. The share of trade with NNTR countries is close to zero throughout and no trade comes in under NNTR after 2005.[9]

Panel (c) plots the share of U.S. imports, identified by the country's regime in 1976. The UTPP countries' share rises from two percent in 1976 to ten percent in 2017, while former-NNTR countries' share rises from zero to nearly 40 percent. Panel (d) shows the decline in tariffs for these groups. Tariffs for former-NNTR countries were substantially above NTR rates and then fell almost to the NTR rates. These drops are discrete and involve a country

Statistics (2017).

[7]For the current list of rates of provision and country sub-codes visit https://www.census.gov/foreign-trade/reference/codes/rp.html and https://www.census.gov/foreign-trade/reference/codes/csc.html, respectively.

[8]The movements in the NNTR rate are caused by composition effects. The statutory NNTR rates have changed little.

[9]In 2022, Belarus and Russia were moved back to the NNTR status.

or set of countries being transitioned to NTR or UTPP. The long-run differences in the tariffs of countries that started with NNTR and NTR rates reflects differences in the composition of goods being imported. In contrast to the large drops in tariffs among countries initially with NNTR rates, the NTR rates fell only about five percent. The rise in trade with countries that transitioned from NNTR to NTR is a key component of the U.S. import growth that we seek to capture.

2.3 Tariff dynamics across vs. within regimes

We now document the key features of trade-policy dynamics, with a focus on comparing tariff changes across versus within regimes. We begin by discussing the dynamics of these regimes themselves. Table 1 reports the persistence of regimes at the country (j) and country-product (jg) levels, and Table 2 lists the five largest (in terms of the number of affected goods) example countries for each type of transition. While all four regimes are highly persistent, there are several transitions away from each of them, indicating that no regime should be viewed as permanent. If our data allowed us to go back further in time, we would observe more transitions. For example, going back to 1949 would add about 20 more country-level transitions from NTR and NNTR and at least two from NNTR to NTR (Yugoslavia and Poland).

We turn now to tariff dynamics. Table 4 reports summary statistics for tariff changes. The typical tariff change is relatively small, with a mean of -0.17 percentage points over one year and -0.73 percentage points over five years. However, tariff changes are much larger when we condition on regime changes. For example, the mean absolute tariff change during a transition into or out of NNTR is almost 30 percentage points. The large tariff changes that occur during regime changes will likely generate longer transitions, which we will allow for in our modeling.

Tariff changes that occur during regime changes are also much more persistent. To illustrate this, we measure the the autocorrelation of tariff changes by estimating the local projection of the horizon-h tariff change on the initial tariff change following Boehm et al.

(2023), but allow for different responses for changes across versus within regimes:

$$\Delta_h \tau_{jgt} = \beta_h^{\tau,within} \Delta_0 \tau_{jgt} \mathbb{1}\{regime_{jgt} = regime_{jgt-1}\} + \beta_h^{\tau,across} \Delta_0 \tau_{jgt} \mathbb{1}\{regime_{jgt} \neq regime_{jgt-1}\} + \delta_{jt} + \delta_{gt} + u_{jgt}. \quad (1)$$

Note we introduce a common set of fixed effects that apply to tariff changes both across versus within regimes. The δ_{gt} fixed effect captures tariff-change variation that is common across countries, e.g. tariff cuts during GATT rounds.[10] We estimate (1) for horizons $h = [-5, 10]$ and cluster standard errors at the jg level. Figure 3(a) plots the paths of $\beta_h^{\tau,within}$ and $\beta_h^{\tau,across}$, as well as estimates from a pooled specification without the interaction terms. Tariff changes across regimes are much more highly autocorrelated than within-regime changes. As we will show, trade responds very differently to transitory tariff changes than to canonical unanticipated once-and for-all reforms. Note also that the within-regime autocorrelation is very similar to the pooled autocorrelation because the sample is mostly comprised of within-regime changes. This indicates that the pooled estimates are poor predictors of the dynamics of across-regime tariff changes.

There is also substantial variation in tariff dynamics within the NTR regime. Specifically, tariff changes following GATT rounds are very different than other within-NTR changes. The Tokyo round, which was negotiated from 1973-1979, involved gradual tariff cuts from 1980 to 1987, while the Uruguay round, which was negotiated from 1986 to 1994, involved gradually phased-in cuts from 1995 to 2000. When we restrict attention to tariff changes during these windows, the five-year tariff is roughly five times the one-year change, while in other periods the one- and five-year changes are similar. The gradual tariff changes during GATT rounds[11] are similar to the tariff "phaseouts" that occur during NTR-to-PTA transitions. In these transitions, the one-year change is about -2.68 percentage points, compared to -4.0 percentage for the five-year change. As we will argue, long negotiations and pre-announced tariff phaseouts also make trade respond very differently to these these kinds of tariff changes than to the canonical unanticipated once-and for-all reform.

[10] In the appendix, we show that excluding these fixed effects magnifies the differences in tariff dynamics across versus within regimes.

[11] We plot the path of tariffs separately for Tokyo and Uruguay rounds in Figure A5

2.4 Case studies: China and Vietnam NTR access

We now describe the dynamics of tariffs on imports from China and Vietnam, which represent the two largest transitions from NNTR to NTR. Figure 2(a) shows how tariffs on China and Vietnam have changed over time relative to NTR tariffs. Specifically, we plot the mean and interquartile range of the ratio of the NTR rate to the tariff on China and Vietnam,

$$\tilde{\tau}_{jgt} = \frac{1 + \tau_{gt}^{\text{NTR}}}{1 + \tau_{jgt}}. \tag{2}$$

This measure is bounded between zero during the embargo and one when these countries gain NTR status. In both cases, we see a long embargo, followed by substantial dispersion during the NNTR regime, which then subsides when each country shifts to NTR.[12] When scaled this way, the transitions from NNTR are most similar to once and for all tariff changes. However, as we have seen, there are transitions across regimes. Moreover, the reasons for the transitions are closely tied to the state of geopolitical relations between each country and the United States.

The tariff changes for China and Vietnam during their transitions from NNTR to NTR were much more persistent than most other tariffs changes, including other across-regime changes. To illustrate this, we estimate the autocorrelation of these specific reforms using the specification,

$$\begin{aligned} \Delta_h \tau_{jgt} = {} & \beta_h^{\tau,china} \Delta_0 \tau_{jgt} \mathbb{1}\{j = china, t = 1980\} \\ & + \beta_h^{\tau,vietnam} \Delta_0 \tau_{jgt} \mathbb{1}\{j = vietnam, t = 2002\} \\ & + \beta_h^{\tau,others} \Delta_0 \tau_{jgt} \mathbb{1}\{other\ j, t\} + \delta_{jt} + \delta_{gt} + u_{jgt}. \end{aligned} \tag{3}$$

Figure 3(b) plots the paths of these estimates. Tariffs on China and Vietnam are extremely persistent compared to tariffs on other countries; statistically, we cannot reject the hypothesis that there are no further changes in the next ten years. Thus, the transitions of China and Vietnam from NNTR to NTR appear to be the closest empirical counterparts to UOAFA

[12] The trade embargo on North Vietnam started in 1964 and was extended to all of Vietnam in 1975. The trade embargo on China was in place from 1950 to 1971.

reforms.

3 Reduced-form evidence on trade dynamics

We now turn to measuring the dynamics of the trade elasticity. Our conceptual object of interest is the path of the cumulative trade elasticity,

$$\varepsilon_h = -\frac{\Delta_h v_{jgt}}{\Delta_h \tau_{jgt}}. \tag{4}$$

We first estimate the dynamics of the trade elasticity to the average tariff, and show that the responses are small and grow little over time. We then condition on regime changes and find larger, more gradual responses. Finally, we estimate trade dynamics for China and Vietnam following their NTR grants and find even larger responses. It is important to emphasize that these estimates are reduced-form responses that reflect agents' beliefs about the stochastic process for tariffs as well as the tariff changes observed in the data. In light of the evidence in section 2, we interpret the differences in trade dynamics for tariff changes within versus across policy regimes as evidence of differences across regimes in the underlying tariff process.

3.1 Trade dynamics across vs. within regimes

As we have seen, most tariff changes, especially those that occur within policy regimes, are transitory. To account for this, we follow the approach of Boehm et al. (2023), which is to estimate (4) using differences in differences,

$$\Delta_h v_{jgt} = -\beta_h^v \Delta_h \tau_{jgt} + \delta_{jt} + \delta_{gt} + u_{jgt}, \tag{5}$$

but using $\Delta_0 \tau_{jgt}$ as an instrument for $\Delta_h \tau_{jgt}$ for $h > 0$. The idea behind this approach is that that tariffs may change several times between period $t-1$ and period $t+h$, and if later changes are correlated with earlier ones, estimating ε_h using ordinary least squares may confound the effects of the former with the effects of the latter. This captures the spirit of local projections as described by Jordà (2005) while correcting for autocorrelation in the tariff process. We include the same set of fixed effects as in section 2, which is standard in the trade-elasticity estimation literature. The country-year fixed effects δ_{jt} control for

aggregate shocks such as exchange-rate movements or aggregate cyclical fluctuations. The good-year fixed effects control for changes in multilateral resistance, such as good-specific fluctuations in U.S. demand as well as multilateral policy changes.[13] Note that by taking differences of the dependent variable, non-time varying variety effects such as trends are eliminated. Again, we provide estimates for horizons $h = [0, 14]$ and cluster standard errors at the jg level.

We first estimate (5) for the full sample, shown by the black line in Figure 4. The short-run trade elasticity, β_0^v, is 2.14, and the long-run elasticity, β_{14}^v, is about 3.48. Note that these pooled estimates not only restrict trade responses to be the same for within- vs. across-regime tariff changes, but also impose the same autocorrelation structure for tariffs in the first stage. We then estimate a version of (5) that allows for different responses for tariff changes across versus within regimes:

$$\begin{aligned}\Delta_h v_{jgt} = &-\beta_h^{v,within} \Delta_h \tau_{jgt} \{regime_{jgt} = regime_{jgt-1}\} \\ &- \beta_h^{v,across} \Delta_h \tau_{jgt} \{regime_{jgt} \neq regime_{jgt-1}\} + \delta_{jt} + \delta_{gt} + u_{jgt}.\end{aligned} \tag{6}$$

The first stage in this extended specification is (1). Thus, we now allow for both different trade responses and tariff processes across versus within reforms. Figure 4 shows these estimates in blue (within) and red (across), respectively. The estimates for within-regime trade elasticities are lower than the pooled estimates and there is little adjustment over time, with a short-run elasticity of 1.61 and a long-run elasticity of 2.11. The across-regime elasticities are noticeably higher and grow more over time, with a short-run value of 3.87 and a long-run value of 5.86. As with our estimates of tariff dynamics, these results indicate that the pooled estimates are not predictive of how trade responds to large reforms that involve regime changes because within-regime tariff changes constitute the vast majority of the sample.

Last, we estimate a version of (5) that compares trade dynamics for China and Vietnam

[13] Our results are robust to the inclusion of more restrictive 1- and 2-digit SITC-country-year fixed effects.

to other countries' trade dynamics:

$$\begin{aligned}\Delta_h v_{jgt} = &- \beta_h^{v,chn+vnm} \Delta_h \tau_{jgt} \{j = chn \vee vnm\} \\ &- \beta_h^{v,other} \Delta_h \tau_{jgt} \{j = other\} + \delta_{jt} + \delta_{gt} + u_{jgt}.\end{aligned} \tag{7}$$

Here, the first-stage is similar to (3), except that we group China and Vietnam together and do not condition on years. The line for China and Vietnam in Figure 4 plots the path of $\beta_h^{v,chn+vnm}$. The short-run elasticity is 3.16, higher than in the pooled sample and for within-regime changes, but lower than for all across-regime changes. The long-run elasticity is 6.91, which is significantly higher than for all across-regime changes. Our interpretation of this is that other regime changes, especially the formation of PTAs, feature a greater degree of anticipation due to long negotiations and tariff phaseouts. As we show in the next section, this increases measured trade elasticities in the short run but also reduces them in the long run. Again, this suggests that the cases of China and Vietnam are better-suited for estimating the how trade responds to UOAFA reforms.

3.2 Case studies: China and Vietnam NTR access

We continue our analysis of trade dynamics by honing in on the effects of the transitions from NNTR to NTR that China and Vietnam underwent in 1980 and 2002, respectively. Leveraging the heterogeneity in tariff changes across goods, we estimate the dynamics of trade following these reforms, as well as in the years leading up to them. Specifically, we follow Alessandria et al. (2025) to estimate the relationship between trade flows and the difference between the NNTR tariff and the NTR tariff, which we refer to as the *NNTR gap*:

$$X_g = \tau_g^{\text{NNTR}} - \tau_g^{\text{NTR}}, \tag{8}$$

where the tariffs are measured in XX. The NNTR gap has been widely used in the literature as a measure of tariff risk, as it represents the tariff hike on goosd g that would occur if a country lost NTR status and moved back to NNTR. However, it is also the tariff reduction that good received when the country originally gained NTR status. This dual meaning complicates the interpretation of the estimates discussed below, but does not preclude them

from being used as inputs to a quantitative analysis, which is the strategy that we ultimately follow.

We use an event-study design to estimate the responses of imports from each country to NTR access:

$$v_{jgt} = \sum_{t'=1974}^{2008} \beta_t^{\text{CH}} \mathbb{1}\{t = t'\} \mathbb{1}\{j = \text{CH}\} X_g^{\text{CH}} + \sum_{t'=1994}^{2017} \beta_t^{\text{VN}} \mathbb{1}\{t = t'\} \mathbb{1}\{j = \text{VN}\} X_g^{\text{VN}} + \delta_{jt} + \delta_{jg} + \delta_{gt} + u_{jgt}. \tag{9}$$

The dependent variable is the log import value of good g from country j in year t. Our coefficients of interest, β_t^j, measure the elasticity of trade with respect to the NNTR gap, relative to the other countries with NTR status and relative to a country-specific base year. We focus on a window from 1974 to 2008 for China to avoid effects from the Great Recession, while for Vietnam we focus on the full sample ending in 2017.[14] As before, we include country-year fixed effects (δ_{jt}) to capture aggregate shocks to exporting countries, country-good fixed effects (δ_{jg}) to capture the average level of exports and time-invariant bilateral trade barriers, and good-year fixed effects (δ_{gt}) to capture good-level U.S. demand shocks and good-specific trade barriers common to all exporters.[15] The elasticities are normalized to zero four years before the tariff change. For China, this is four years after the embargo ends, while, for Vietnam, this is three years after the embargo ends. [16]

Figure 5(b) shows the estimates of β_t^{CH} and β_t^{VN}. There are four key observations. First, both countries have similar long-run NNTR-gap elasticities that are substantially higher than the average long-run trade elasticities we previously estimated. From trough-to-peak, the elasticity grows by about 12 log points for Vietnam and 10 for China. Second, the Chinese response, on impact, was about half as big as the Vietnamese response. Third,

[14] As a major U.S. trade partner at that time, the Great Recession had a significant impact on U.S. imports from China. Trade with Vietnam was still in the very early stages and was not significantly impacted.

[15] This fixed-effects structure has been used frequently in the literature on China's NTR access, including by Pierce and Schott (2016) and Handley and Limão (2017).

[16] Note the categorical variable distinguishes imports from China in 1980, Vietnam in 2002, and all other observations. Hence, in the case of the former two, only varieties with non-zero trade in the year of and the year before the liberalization are used to identify the trade elasticities, explaining their relatively large confidence intervals.

the two countries have different trade dynamics in the lead up to the tariff change. Trade with Vietnam in high-gap goods was growing faster than low-gap goods while with China these different goods were growing at about the same pace. Fourth, Chinese exports grew much slower than exports from Vietnam; it took more than twelve years to reach China's long-run level, whereas Vietnam's adjustment was more than 90 percent complete after four years (note that this is precisely the length of the period in which the Vietnamese reform remained conditional).

The large response in trade to large, persistent country- and good-specific changes in tariffs is consistent with the idea that trade responds more to more persistent tariff change. However, the fact that the tariff change was ex-post more persistent does not tell us about the expected path of tariffs at each moment in time, which depended on complex geopolitical risks. Indeed, through the lens of the model we present next, the differences in the short-run response and speed of convergence across countries suggests China's NTR access was initially less credible and took longer to gain credibility, while Vietnam's access was more highly anticipated and gained credibility more quickly. We now turn to matching these dynamics with in a structural model with country-specific trade-poluicy expectations that leverages the precise form on uncertainty faced by countries transiting between NNTR and NTR tariffs.

3.3 Robustness

1. Together vs Individual
2. Different gaps
3. HS/TSUSA vs SITC
4. Short-run based on first differences.
5. Long-run first differences.

4 Model

To illustrate how policy dynamics complicate the measurement of trade dynamics, we use a model in which firms make forward-looking export participation decisions to show how trade

responds to uncertain policy. Our model is a partial-equilibrium version of Alessandria et al. (2021) in which a continuum of heterogeneous firms choose to start and stop exporting in the face of idiosyncratic technology shocks and aggregate shocks to trade policy. The model has two key ingredients that generate gradual adjustment to policy: sunk exporting costs and export capacity that grows stochastically with a firm's tenure as an exporter. The first ingredient makes firms consider the entire future path of trade policy, not just its current state (Das et al., 2007; Alessandria and Choi, 2007). The second ingredient generates exporter life-cycle dynamics (Fitzgerald et al., 2016; Alessandria et al., 2021; Steinberg, 2023), which captures the behavior of marginal exporters and non-exporters. The model thus allows for potentially large differences between short- and long-run trade elasticities.

Production and demand. Firms use labor, ℓ_t, to produce output according to constant returns to scale technologies:

$$y_t = z_t \ell_t. \tag{10}$$

A firm's productivity, z_t, evolves over time according to an AR(1) process in logs:

$$\log z_t = \rho_z \log z_{t-1} + \sigma_z \varepsilon_t, \tag{11}$$

where ε_t is i.i.d. across firms and time. Foreign demand for a firm's product is a downward-sloping function of the price the firm charges, p_t, and the import tariff, τ_t:

$$d_t(p_t, \tau_t) = (p_t \tau_t)^{-\theta}, \tag{12}$$

where θ is the price elasticity of demand.

Trade costs. There are three types of trade costs: import tariffs; iceberg-type variable trade costs; and fixed trade costs. Import tariffs apply to all firms equally and are, at this stage, allowed to follow an arbitrary stochastic process. We index model objects by the time period, t, which is a parsimonious way to encode firms' expectations about future tariffs. Variable trade costs can take three values, $\infty > \xi_H > \xi_L$, that encode a firm's status as an exporter. A firm with $\xi = \infty$ is a non-exporter, a firm with ξ_H is a low-capacity exporter, and

a firm with $\xi = \xi_L$ is a high-capacity exporter. In order to begin exporting, a non-exporter must pay a fixed cost f_0. In order to continue exporting, a low- or high-capacity exporter must pay a fixed cost f_1. Thus, the fixed cost of exporting can be expressed as a function, $f(\xi)$, of the variable cost, where $f(\infty) = f_0$ and $f(\xi_H) = f(\xi_L) = f_1$.

Exporter life cycles. Firms are born as non-exporters with $\xi = \infty$. In each period, firms die with probability $1 - \delta(z) = \max\left[0, \min\left(e^{-\delta_0 z} + \delta_1, 1\right)\right]$, in which case they are replaced by newborn firms with productivities drawn from the ergodic distribution associated with the process (11). When a non-exporter chooses to enter the export market, its variable trade costs falls to ξ_H in the next period. When a low-capacity exporter chooses to continue exporting, its variable exporting cost falls to ξ_L in the next period with probability $1 - \rho_\xi$. Symmetrically, when a high-capacity exporter chooses to continue exporting, the probability of switching back to ξ_H is also ρ_ξ.

Firm's problem. The firm's state variables are its productivity, z, and its variable trade cost, ξ. The firm's problem has a static component and a dynamic component. The former entails choosing a price to maximize the profits from exporting in the current period:

$$\pi_t(z_t, \xi_t, \tau_t) = \begin{cases} \max_p \left\{ p d_t(p, \tau_t) - w_t \frac{\xi_t d_t(p, \tau_t)}{z_t} \right\} & \text{if } \xi \in \{\xi_L, \xi_H\} \\ 0 & \text{if } \xi = \infty \end{cases} \tag{13}$$

The latter entails deciding whether to export in the next period. The value of a firm that chooses to export in $t+1$ is

$$V_t^1(z_t, \xi_t, \tau_t) = -f(\xi_t) + \frac{\lambda}{1+r} \mathbb{E}_{z,\xi,\tau} \left[V_{t+1}(z_{t+1}, \xi_{t+1}, \tau_{t+1}) \right], \tag{14}$$

and the value of a firm that chooses not to export is

$$V_t^0(z_t, \xi_t, \tau_t) = \frac{\lambda}{1+r} \mathbb{E}_{z,\tau} \left[V_{t+1}(z_{t+1}, \infty, \tau_{t+1}) \right], \tag{15}$$

where

$$V_t(z_t, \xi_t, \tau_t) = \pi_t(z_t, \xi_t, \tau_t) + \max\{V_t^0(z_t, \xi_t, \tau_t), V_t^1(z_t, \xi_t, \tau_t)\}. \tag{16}$$

The solution to the export participation problem is characterized by a break-even level of productivity, $\bar{z}_t(\xi)$, such that $V_t^0(\bar{z}_t(\xi), \xi, \tau_t) = V_t^1(\bar{z}_t(\xi), \xi, \tau_t)$. Importantly, $\bar{z}_t(\xi)$ depends on firms' expectations about future tariffs, not just the current tariff rate. We will consider several processes for tariffs.

Aggregation. Aggregate exports are given by

$$X_t = \sum_{\xi \in \{\xi_L, \xi_H\}} \int_z p_t(z, \xi, \tau_t) d_t(p_t(z, \xi, \tau_t), \tau_t) \, \mathrm{d}\varphi_t(z, \xi), \tag{17}$$

where $\varphi_t(z, \xi)$ denotes the joint distribution of firms indexed by their productivity and variable trade cost.

Parameters. For the numerical examples in the next section, we follow Alessandria et al. (2021). We assign several values directly from that paper: $\theta = 4$, $\rho_z = 0.65$, $\sigma_z = 1.32$, $\rho_\xi = 0.92$, $\delta_0 = 21$, $\delta_1 = 0.02$. We set the interest rate, r, to 4 percent. We normalize the low variable trade cost, ξ_L, to one and choose values for the fixed costs, f_0 and f_1, and the high variable trade cost, ξ_H, so that the model's free-trade steady state reproduces three moments from Alessandria et al. (2021): (i) 22.3 percent of firms export; (ii) 17 percent of exporters exit; and (iii) new exporters sell half as much as incumbent exporters on average. We discuss our parameterization of the quantitative model for the case studies of China and Vietnam in section 6.1.

5 Illustrative examples

We now use the model to illustrate some issues in estimating the trade response to a change in tariffs. We conduct three kinds of hypothetical reforms in the model: unanticipated, once and for all reforms, stochastic reforms, and changes to beliefs about the policy's stochastic process. Each reform highlights a different issue in measuring trade adjustment to changes in trade policy that can be related to the changes in trade policy regimes. We also consider how the initial conditions and the window of observation influence our measures of the trade response.

When tariffs follow a Markov process, the anticipation of future tariff changes and persistence of past changes simultaneously affect trade volumes. Consider a tariff process with two states, $\tau_L = 0$ and $\tau_H = 30$ percent, and a symmetric probability of switching states, $1 - \omega$. We conduct the same reform—switching from the high-tariff state to the low-tariff state—for values of ω ranging from 0.5 to 1. When $\omega = 0.5$, tariffs are i.i.d. across time, and the current tariff contains no information about future tariffs. When $\omega = 1$, tariffs are constant and the tariff cut is permanent. In between these extremes, tariffs are persistent but firms know there is a chance of switching states. In all versions of this experiment, the economy begins in the risky steady state with high tariffs, switches to the low-tariff state in period $t = 0$, and remains there forever.

It is useful to summarize the future path of tariffs by the *mean discounted expected tariff*,

$$\bar{\tau}^E_{t+1} = (1 - \beta)^{-1} E_t \sum_{s=t+1}^{\infty} \beta^{s-t-1} \tau_s. \tag{18}$$

In Figure 6, we show how the path of $\bar{\tau}^E_{t+1}$ changes as we vary the probability of remaining in the same state.

5.1 Canonical reform: unanticipated, immediate, and permanent

As a benchmark, we consider the canonical trade reform found in virtually all quantitative trade studies: an unanticipated, immediate, and permanent reform. This is the tariff process with $\omega \to 1$. Figure 6 shows the results of this reform in our model (line $\omega = 1.0$). In period zero, the distribution of firms over trade costs and trade participation is predetermined, so trade only responds on the intensive margin, and the short-run trade elasticity is equal to the CES demand elasticity: $\varepsilon_0 = \theta$. We refer to this as the *short-run canonical trade elasticity*. In period one, export participation begins to rise, and converges to its new long-run value, which is about four log points higher than before the reform (Figure 6(c)). Trade follows a similar trajectory, and the long-run trade elasticity, ε_∞, is about 17.5 (Figure 6(e)). We refer to this as the *long-run canonical trade elasticity*. It captures the long-run change in trade relative to the long-run change in tariffs.

5.2 Stochastic reforms

When ω is not one, the mean discounted expected tariff differs from the current tariff and thus the measured trade elasticities will deviate from their canonical values. The less persistent are tariffs, the smaller is the change in the mean discounted tariff when the tariff changes and the lower is the long-run trade elasticity to the tariff change. This result is driven by two effects. The first is an anticipation effect. When switching from the high-tariff state to the low-tariff state is more likely, exporting is more valuable in the high-tariff state, which increases export participation and aggregate exports in the initial risky steady state. To highlight this effect, we plot changes in export participation and trade in Figures 6(c) and 6(d), relative to the pre-reform period in the model in which tariffs are believed to be constant (i.e., one where $\omega \to 1$). When tariffs are less persistent, export participation and trade are higher before the reform occurs because firms believe it is more likely to occur in the future. The second effect depresses trade after the tariff cut. When switching states is more likely, the reform is expected to be reversed sooner, depressing export participation and aggregate trade after the reform.

When tariffs are less persistent, the pre-reform level of trade is higher and post-reform trade growth is lower. Both of these effects push the measured long-run trade elasticity below the canonical long-run elasticity. The i.i.d. case is particularly instructive. In this case, the current tariff state does not influence the mean discounted expected tariff and thus the value of exporting in the future is constant. Consequently, export participation does not react to tariff shocks at all, and the trade elasticity at all horizons is equal to the demand elasticity, θ.

5.3 Shocks to expectations

In the previous section, when tariffs changed, the mean, discounted expected tariff could also change. Here, we consider changes in the mean, discounted expected tariff that may not coincide with a change in tariffs. To do so, we shift from a two- to four-state Markov transition matrix that allows for different transition probabilities from the same tariff levels. Specifically, the first two states have a tariff of 30 percent and the last two states have a zero tariff. The persistence of the first and fourth states are close to one while the second

and third states are less persistent and the transition between the 2 and 3 states is much more likely than those states to either the first or fourth. In this way we can separately study changes in mean discounted expected tariffs from changes in the tariff. Our four state transition matrix is

$$\Omega = \begin{bmatrix} 1-3\epsilon & \epsilon & \epsilon & \epsilon \\ \epsilon & \omega-\epsilon & 1-\epsilon-\omega & \epsilon \\ \epsilon & 1-\epsilon-\omega & \omega-\epsilon & \epsilon \\ \epsilon & \epsilon & \epsilon & 1-3\epsilon \end{bmatrix}. \tag{19}$$

We start in state 1 with high tariffs that are expected to be permanent. As before, we assume we have been in this state for many years. We assume that $\omega = 0.8$ and $\epsilon \approx 0$. In Figure 7, we plot the response to four reforms. Our first reform is a transition from state one to state four at $t = 1$. This is the canonical unanticipated once and for all reform we have already discussed.

Our second reform is a transition from state one to state two. This keeps the tariff constant but lowers the mean discounted expected tariff to 15.8 percent. We assume the shift occurs in period $t = -1$. Trade jumps in period 0 and grows persistently to about 1.58 log points above the initial level. Without a change in the observed tariff, we can not construct an h-on-1 trade elasticity.

Our third reform is from state 1 to state 3. This removes the tariff, but as the change is less persistent than our canonical reform the mean discounted expected tariff only falls to 14.2 percentage points. We get a large and gradual increase in trade. The short-run trade elasticity is equal to 4 and the long-run is equal to 13.5 as trade grows by 3.54 log points. The much larger long run response compared to our second reform, which reduces the expected tariff by 10 percent less, arises from the strong effect of actually having a lower current tariff.

Finally, we consider a transition from state 1 to state 2 followed by a transition to state 3. This uncertain two step reform has the same long-run change in trade as in our third reform, but our measured h-on-1 trade elasticity is a bit larger in the first period, about 7.5. The larger response with the change in tariff occurs because the firms have made investments to

increase market access in the period before the tariff falls and thus trade is responding to the past and current tariff shocks.

5.4 Initial Conditions

In our prior analysis, our model economy starts with a constant transition matrix but a sequence of high tariff draws. This form of stationarity does not capture the transition from tariffs in the U.S. nor the details from Vietnam or China, where the reform followed a long embargo. To capture the role of initial conditions, we start from an embargo and then consider a permanent change. To align with our empirical approach, we do this for two industries that go to different tariff levels of zero and 30 percent.

Figure 8 depicts trade dynamics starting from the period following the embargo's end. We also estimate the trade elasticity from these two industries using a similar panel approach to our estimate of the gap elasticity from (9),

$$v_{gt} = \sum_{t'=1}^{25} \beta_t \mathbb{1}\{t = t'\} \log(\tau_g) + \delta_t + u_{gt}. \tag{20}$$

The good that goes to zero tariffs starts with more trade and exporters and these gaps grow over time. Measured in terms of the trade elasticity, we find a very large trade elasticity from comparing the differences in trade by tariffs of -13.8. Over time, we see that this elasticity gradually falls to about -17.5, which is about the long run trade elasticity we found in the model for an unanticipated once and for all permanent liberalization. Importantly, five years after lifting the embargo we have only closed about two-thirds of the gap between the first year differences and the long-run differences across industries. We have seen that the faster export growth of low tariff goods following the lifting of the embargo is a feature of the trade dynamics in China and Vietnam prior to gaining NTR tariffs. This experiment makes clear that both initial conditions and the length of the observation window may bias empirical estimates of the trade elasticity downward.

6 Quantitative analysis

We take our model to the data—which are contaminated by anticipation of and uncertainty about tariff changes and the gradual effects of past reforms—and use it to measure how trade would respond in the absence of these expectational biases and initial conditions. Our strategy is to discipline the model's technological parameters and the stochastic process for trade policy with empirical evidence from two landmark trade reforms: the United States granting NTR status to China in 1980 and Vietnam in 2002. We use the model to simulate unanticipated, permanent reforms.

The Chinese and Vietnamese reforms have three key advantages. First, the distribution of tariffs and uncertainty are well-understood owing to the structure of trade policy. Second, given the move from the embargo we know the initial conditions in each country and in each sector. Third, the countries receiving the reforms become very large trading partners.

6.1 Environment and calibration

In what follows, we describe the model of firms in China and Vietnam. The structure of the economy is the same in the two countries, but parameter values and timing differ. For clarity, we omit country labels. To exploit the heterogeneity in trade dynamics across products documented above, we consider a multi-product version of the model in section 4. There are G products, which correspond to the 5-digit SITC products in our data. In each product $g = 1, \ldots, G$, a continuum of firms sell differentiated varieties. The technological parameters—the productivity process and the variable and fixed costs of exporting—are the same across products.[17] Tariffs, $\tau_{gt}(s)$, differ across products and across trade-policy regimes, $s \in \{\text{NNTR}, \text{NTR}\}$. The tariff regime, which is an aggregate state, follows a time-varying Markov process with transition probabilities

$$\Omega_t = \begin{bmatrix} \omega_t^{\text{NNTR}} & 1 - \omega_t^{\text{NNTR}} \\ 1 - \omega_t^{\text{NTR}} & \omega_t^{\text{NTR}} \end{bmatrix}, \tag{21}$$

[17] We have explored heterogeneity across products in these parameters as in Alessandria et al. (2025). While this works well for China, there are too few firms in many sectors in Vietnam to reliably estimate exporter-dynamics.

where ω_t^s is the persistence of state s in year t. Our aim is to estimate a path $\{\omega_t^{\text{NNTR}}, \omega_t^{\text{NTR}}\}_{t=0}^{\infty}$ that is consistent with the dynamics of trade flows observed for each country.

Our calibration strategy builds on Alessandria et al. (2025). First, we set several parameters externally. We use the same values for the interest rate, r, the survival rate, δ_0 and δ_1, and the iceberg-cost transition probability, ρ_ξ, from section 5. However, we set the demand elasticity, θ, to Soderbery (2015)'s average estimate for U.S. imports from China, which is 3.13. Soderbery's estimate comes from price and quantity variation and thus is closely related to our demand elasticity parameter.[18] We use the model to explore how the measured short-run trade elasticity we recover depends on the nature of changes in trade policy expectations as we discussed in section 5.3.

Second, we calibrate the other technological parameters, σ_z, f_0, f_1, and ξ_H, to match four moments computed from firm-level panel data: the export participation rate, the export exit rate, the incumbent size premium (the average sales of incumbent exporters relative to the average for new exporters), and the coefficient of variation of log exports. In both the model and the data, we measure these moments many years after NTR status has been granted (2004–2007 for China and 2010–2014 for Vietnam).[19] The entry cost, f_0, is mainly identified by the export participation rate. The continuation cost, f_1, is mainly identified by the exit rate. The initial iceberg trade cost, ξ_H, is mainly identified by the incumbent premium. Productivity dispersion, σ_z, is mainly identified by the coefficient of variation of log exports. Table 5 lists the target moments and calibrated values of the technological parameters for each country.

Third, we calibrate the trade-policy persistence parameters, ω_t^{NNTR} and ω_t^{NTR}, to match the NNTR-gap elasticity dynamics described in section 3.2. ω_t^{NNTR} is identified by the dynamics during the period before the country was granted NTR status. When the NNTR-gap elasticity grows more negative during this period, i.e., exports of high-gap goods grow faster than exports of low-gap goods, it indicates an increase in the likelihood of gaining NTR status. ω^{NTR} is identified by the NNTR-gap elasticity dynamics after the NTR grant. A decrease in this elasticity during this period indicates an increase in the likelihood of

[18] We also explored estimating θ directly in our system and recovered a similar estimate.
[19] The appendix contains more details about the firm-level datasets.

retaining NTR status. The main difference between our strategy and that of Alessandria et al. (2025) is that we allow ω_t^{NNTR} to vary over time, which allows us to capture changes in the probability of the NTR grant occurring in advance of the policy. Figure 9(c) shows the fit of the model against these target moments.[20]

Fourth, we start the model from the embargo so that there are no exports to the United States from either country. We assume that firms do not expect the embargo to be lifted. Changes in the probabilities of gaining and losing NTR status are believed to be permanent. Firms are surprised when these probabilities change.

Figure 9(b) plots the calibrated regime transition probabilities, $1 - \omega_t^s$. For China, the probability of gaining NTR status in the 1970s was low and did not change materially. When China gained NTR status in 1980, the pattern of trade growth is consistent with the view that this status was unlikely to be maintained, and only started to become credible in the mid-1980s. Alessandria et al. (2025) discuss these considerations in detail. For Vietnam, the initial probability of gaining NTR status was similar to China's, but during the late 1990s and early 2000s this probability more than tripled. The model interprets the fast growth of high-gap exports relative to low-gap exports in advance of the Vietnamese NTR grant as evidence of an increase in the expectation that this reform would occur. After Vietnam gained NTR status in 2002, the fast growth in trade is consistent with the probability of losing this status falling faster than in China. As the long-run changes in the NNTR-gap elasticity are similar in the two cases, we estimate similar initial and final transition probabilities.

While we estimate our model to match the trade response in a limited window, our model yields a forecast of the trade beyond that window. For both countries, we see that the trade elasticity continues to increase by 1 percentage point in the five years following our estimation period. The continued trade growth so late in the transition from the tariff cut reflects the time it takes for the reform to become credible.

6.2 Trade response to canonical reforms

We use the calibrated model to measure how trade responds to the canonical tariff reform that features neither anticipation nor uncertainty, is not affected by initial conditions, and for

[20] We HP-filter the NNTR-gap coefficients to smooth out temporary spikes. The calibrated model exactly matches this smoothed series.

which we can observe the entire transition. For each country, we first solve for the stationary distribution under the assumption that NNTR tariffs are permanent and the country has been in this state for a long time. We then simulate an unanticipated, once and for all reduction to NTR tariffs for each country and allow the economy to converge to its new steady state. Importantly, the realized tariff paths in these experiments are exactly the same as in the baseline model with tariff uncertainty, so the differences in trade dynamics are attributable only to differences in expectations and initial conditions.

Figure 9(c) compares the NNTR-gap elasticity dynamics in the canonical NTR reforms to the baseline dynamics. There are three key differences. First, the pre-NTR gap elasticities in the baseline model (which range from about -8 to -11) are smaller than the gap elasticities in the NNTR steady state (about -14). This is largely because firms in the baseline model have a greater incentive to begin exporting due to the chance of gaining NTR status, particularly firms that produce high-gap goods which have the most to gain from such a reform. Moreover, in the data we are not starting from a steady state but on the transition from the embargo, which we have shown biases our long-run estimates downward.

Second, the gap elasticities converge more slowly after the baseline NTR reforms than after the canonical reforms. This is because the probability of keeping NTR status is initially low in both countries, and rises gradually over time. It is particularly noticeable in the case of China, where the gap elasticity stalls for several years after the baseline NTR reform and only starts to rise again several years later.[21] This implies that trade should continue to grow beyond our estimation window for both countries. In our canonical reform there is very little continued growth outside our estimation period.

Third, in the case of Vietnam, the gap elasticity begins to rise several years before the baseline NTR reform, whereas there is no change in advance of the canonical reform. This is an anticipatory effect driven by the increase in the probability of Vietnam gaining NTR status during 1999–2001 shown in Figure 9(b); note there is no anticipatory growth in the case of China, where the likelihood of gaining NTR status was more or less constant throughout the 1970s.

[21] See Alessandria et al. (2025) for a detailed discussion of the geopolitical background of the "stalled" period.

Finally, we calculate the speed of adjustment of the canonical reform. Specifically, along the transition, we measure the gap final trade and current trade in each period relative to the long-run change,

$$\varepsilon_h = \frac{\log V_\infty - \log V_{t+h}}{\log V_\infty - \log V_{t-1}}, \tag{22}$$

where V_∞ is the long-run level of trade in the new steady state with no uncertainty. We then plot this gap in Figure 9(d) separately for China and Vietnam on a log scale. As we have seen, the model predicts very strong growth in the first two years owing to the separate intensive margin and delayed extensive margin response from the one period export lag. After that initial burst, we find that half the distance to the long-run closes about three to five years, with faster convergence in Vietnam than China.

7 Conclusions

We estimate that the long-run trade response to a surprise and perfectly credibly once and for all tariff change is about 15 times the tariff change. We argue that this large response is hard to find in the data owing to several features of the stochastic properties of trade policy and forward-looking decisions of firms related to anticipation, uncertainty, and initial conditions that are not captured in existing empirical methods. We undertake a model-based analysis of trade and trade policy dynamics that leverages key elements of the structure of U.S. trade policy and its relationships with two specific major trade partners - China and Vietnam - to recover these trade elasticities. These reforms were incredibly large - going from embargo to normal trade relation rates - and characterized by a high probability of reversal owing to the unique geopolitical relations between these countries and the U.S.

Our analysis also allows us to also construct the dynamic path of adjustment to an unanticipated once and for all permanent tariff change. We find the short-run elasticity of about 3 and that half of the gap between the current level of trade and long-run is closed in three to five years. Our results imply that there is a mismatch between empirical estimates of trade elasticities and the structural parameters used in quantitative analyses. The vast majority of the tariff variation in the data is highly transitory, so the long-run elasticities identified using this variation are much lower than the elasticities that are appropriate to

use in models of permanent, unanticipated trade reforms. Progress needs to be made on two fronts: the empirical literature should develop methods to estimate trade elasticities that control for anticipation and uncertainty; and the quantitative literature should develop calibration strategies that explicitly account for the stochastic processes that tariffs follow in the data.

If the aim of trade policy is to foster deeper integration, then it is important to find ways to make trade policy more credible. One possibility would be to adopt Milton Friedman's suggestion of a constitutional amendment that sets import tariffs to zero (Friedman and Friedman, 1980), much like the proscription on states taxing interstate commerce or imposing export taxes.[22]

[22]The commerce clause (Article 1, Section 8, Clause 3) gives Congress the power to regulate commerce between states. The export clause (Article 1, Section 9, Clause 5) states "No tax or duty shall be laid on articles exported from any state."

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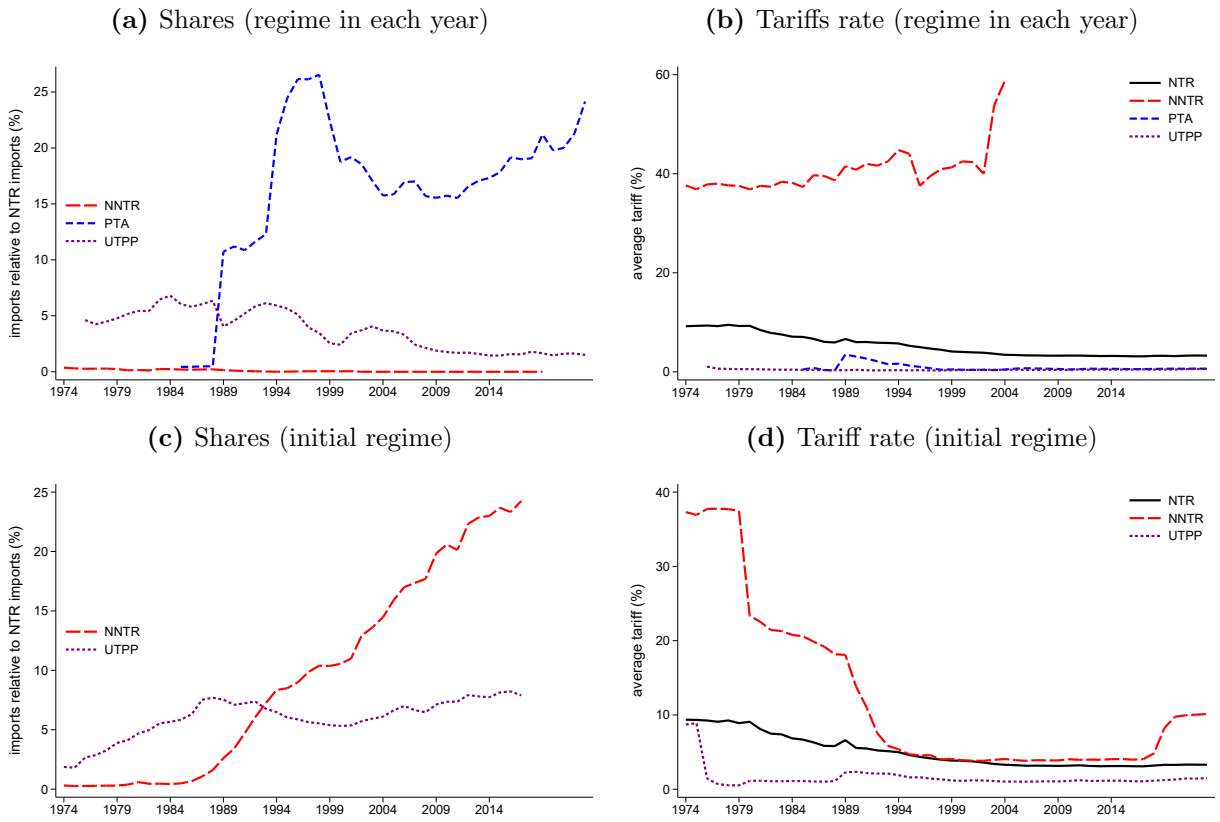
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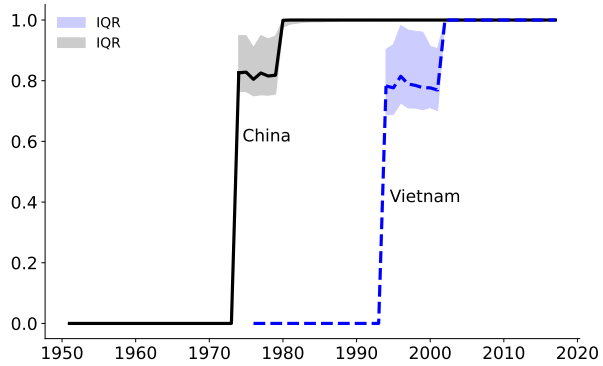
Figure 1: Import shares and tariffs by policy regimes



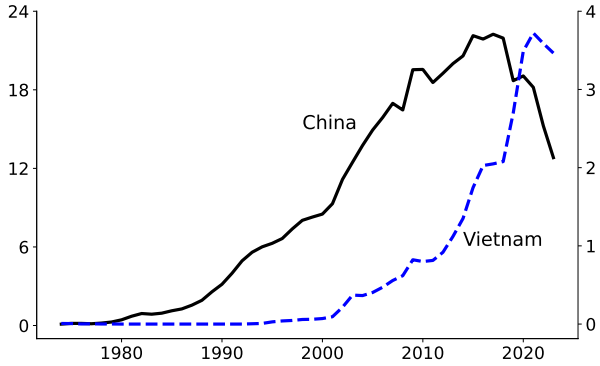
Notes: Panel (a) plots the import share of each of our four trade policy regimes described in 2, after dropping unclassified variety-years. Panel (b) plots the average tariff rate of across products by the regime in that year. Panels (c) and (d) are the respective analogues of panels (a) and (b) where the regime for each variety is fixed across the years as the modal regime in 1970s.

Figure 2: Trade and tariffs in China and Vietnam

(a) Tariff dynamics



(b) Share of US imports (percent)



(c) Extensive margin dynamics

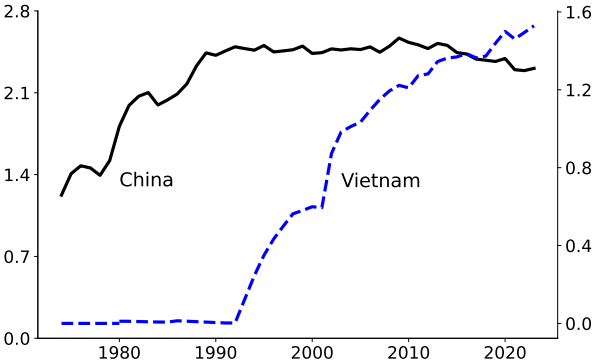
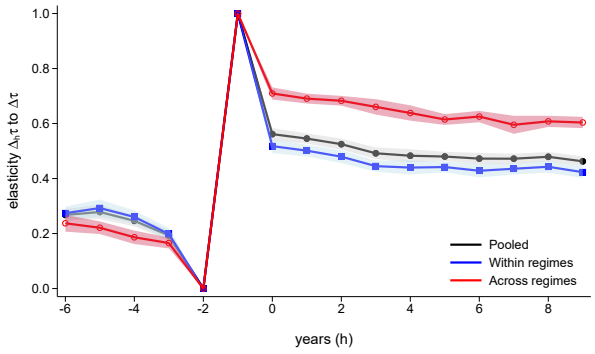


Figure 3: Autocorrelation of tariff changes

(a) Across and Within Regimes



(b) China and Vietnam Regime Change

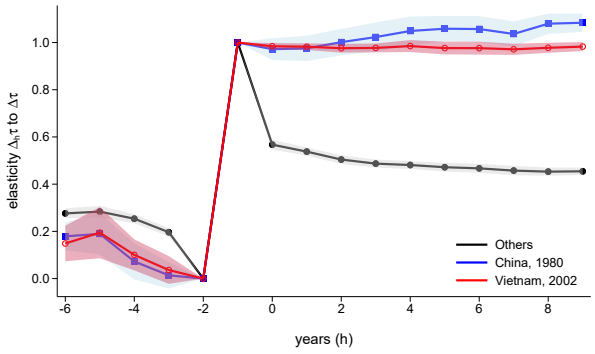


Figure 4: Trade elasticity within and across regimes

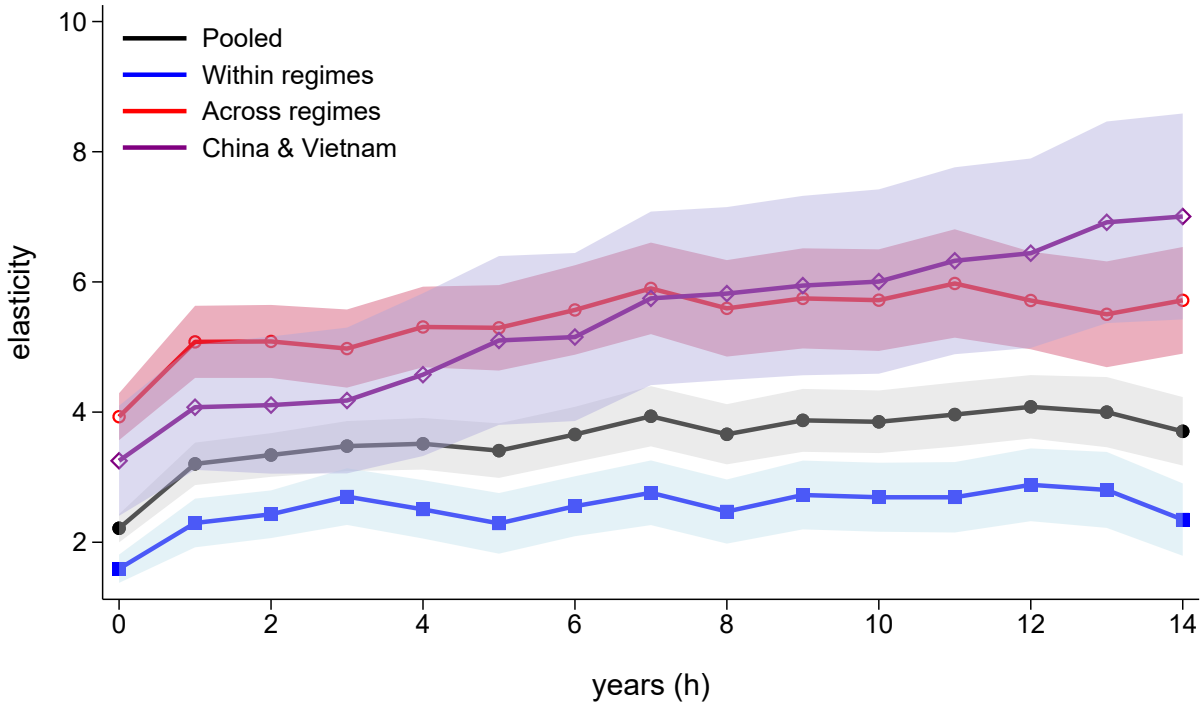
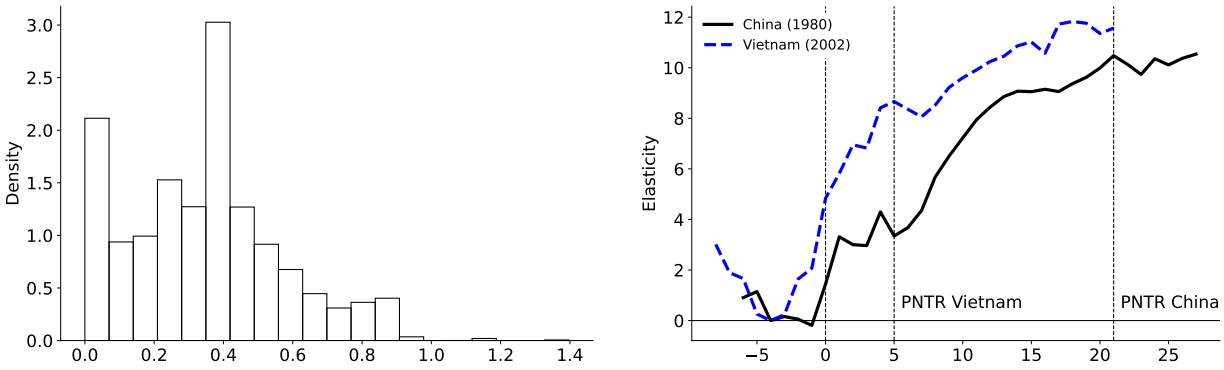


Figure 5: The NNTR liberalizations of China (1980) and Vietnam (2001)

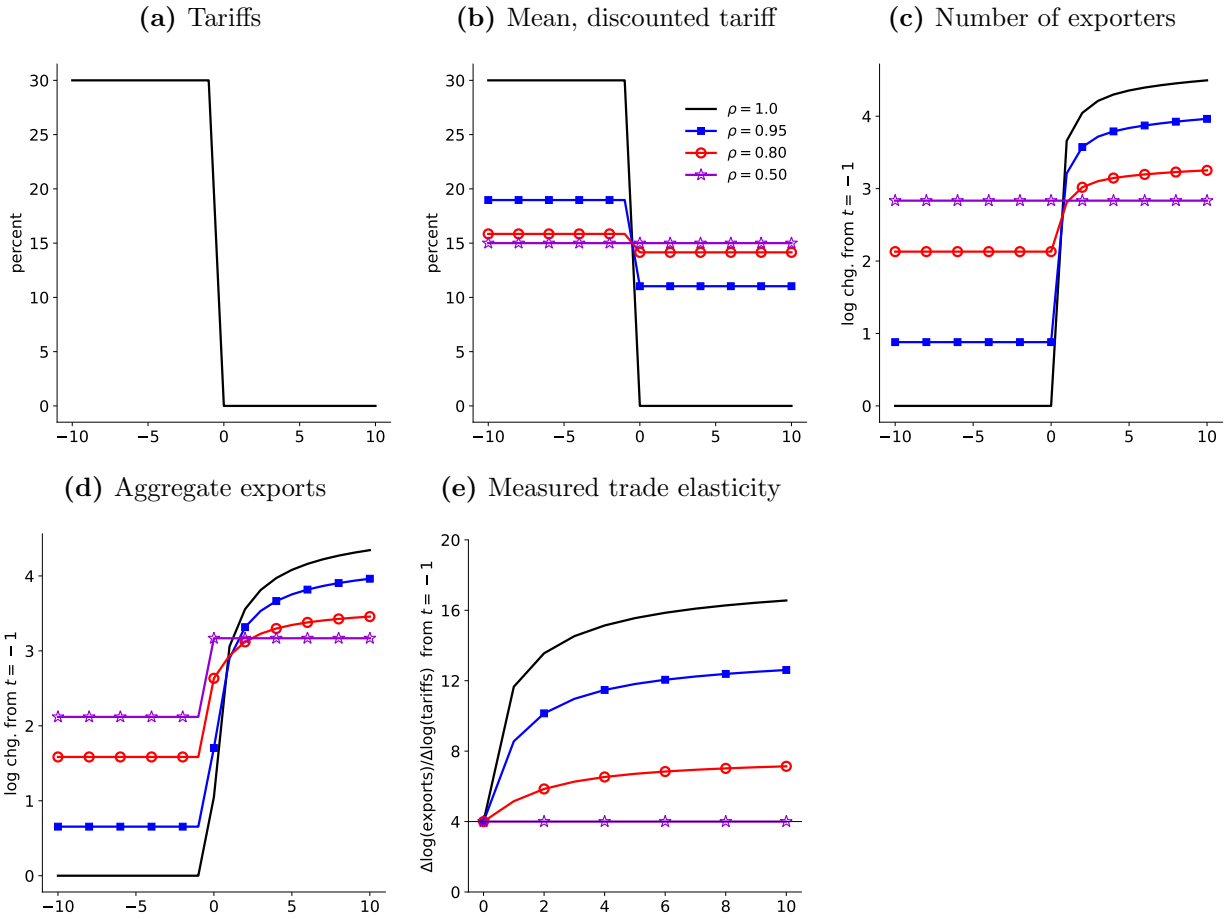
(a) Gap between NNTR and NTR

(b) NNTR gap elasticity



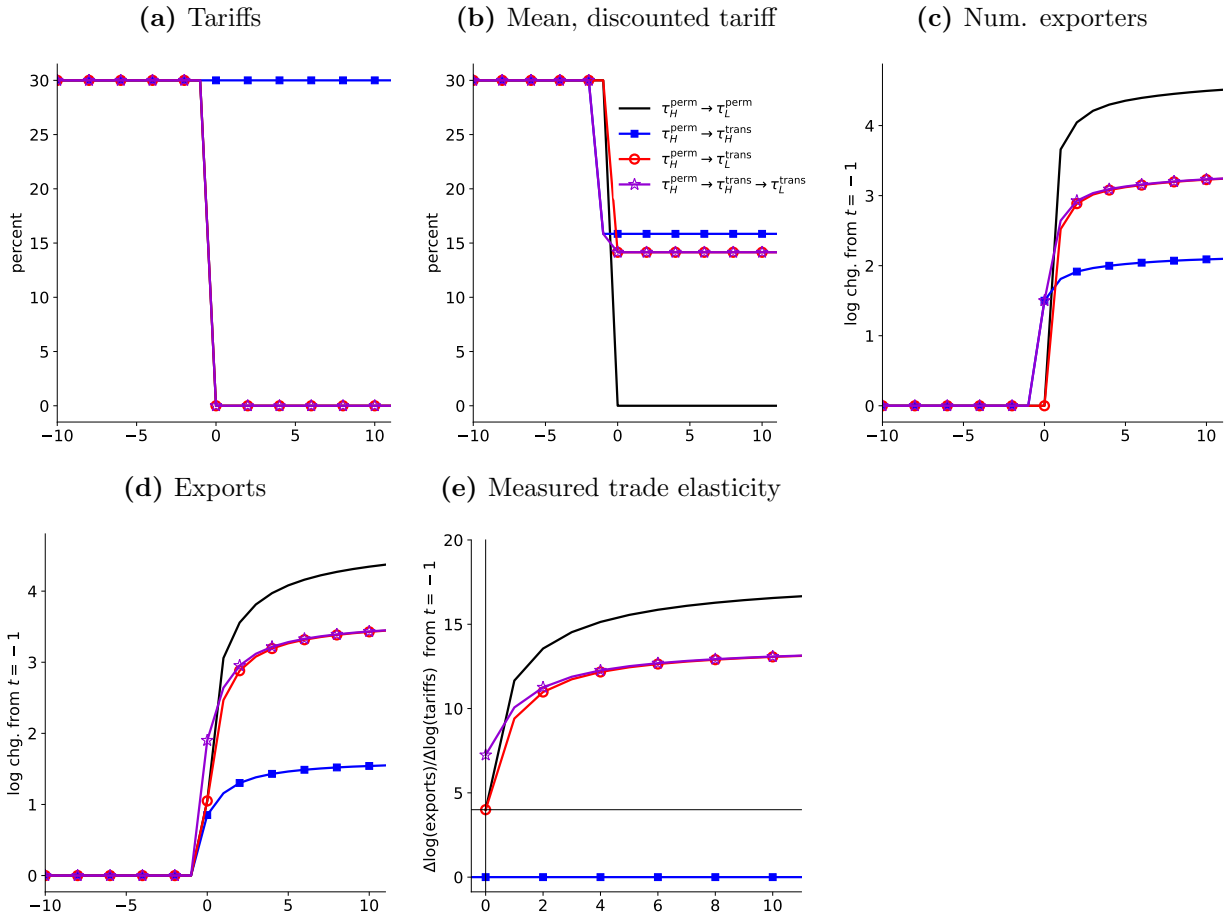
Notes: Panel (a) plots the distribution of the gap between the NNTR and the NTR rate for 1999. Panel (b) plots the elasticity coefficients from (9). The standard errors that construct the 95 percent confidence intervals are clustered at the jg level.

Figure 6: Model responses to Markov reforms



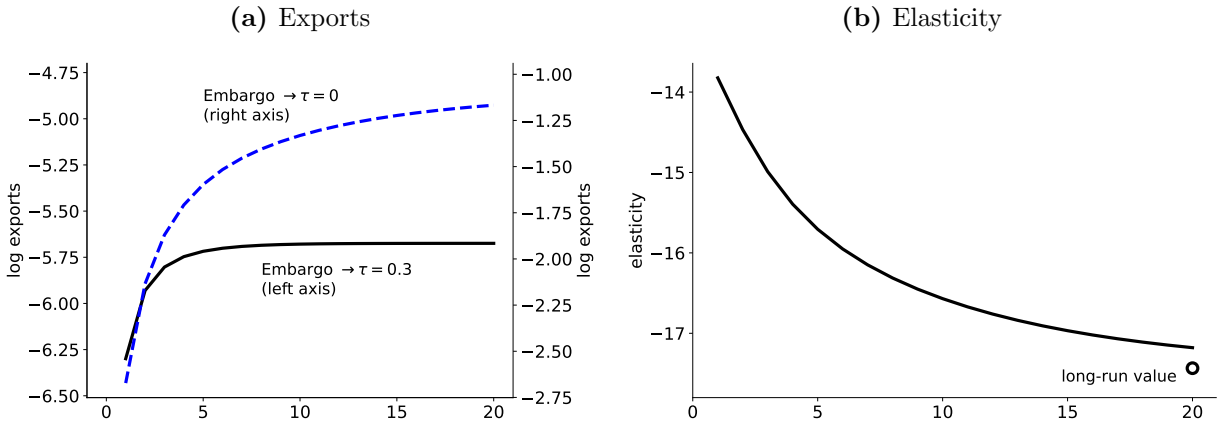
Notes: The figure plots the transition dynamics from the trade reform experiments described in section 5.2, in which the persistence of tariffs (ρ) varies. In panel (e), the measured trade elasticity is defined in (4). In panels (c) and (d), outcomes are measured relative to period $t = -1$ in the calibration with perfectly persistent tariffs ($\rho = 1$).

Figure 7: Model responses to change in persistence



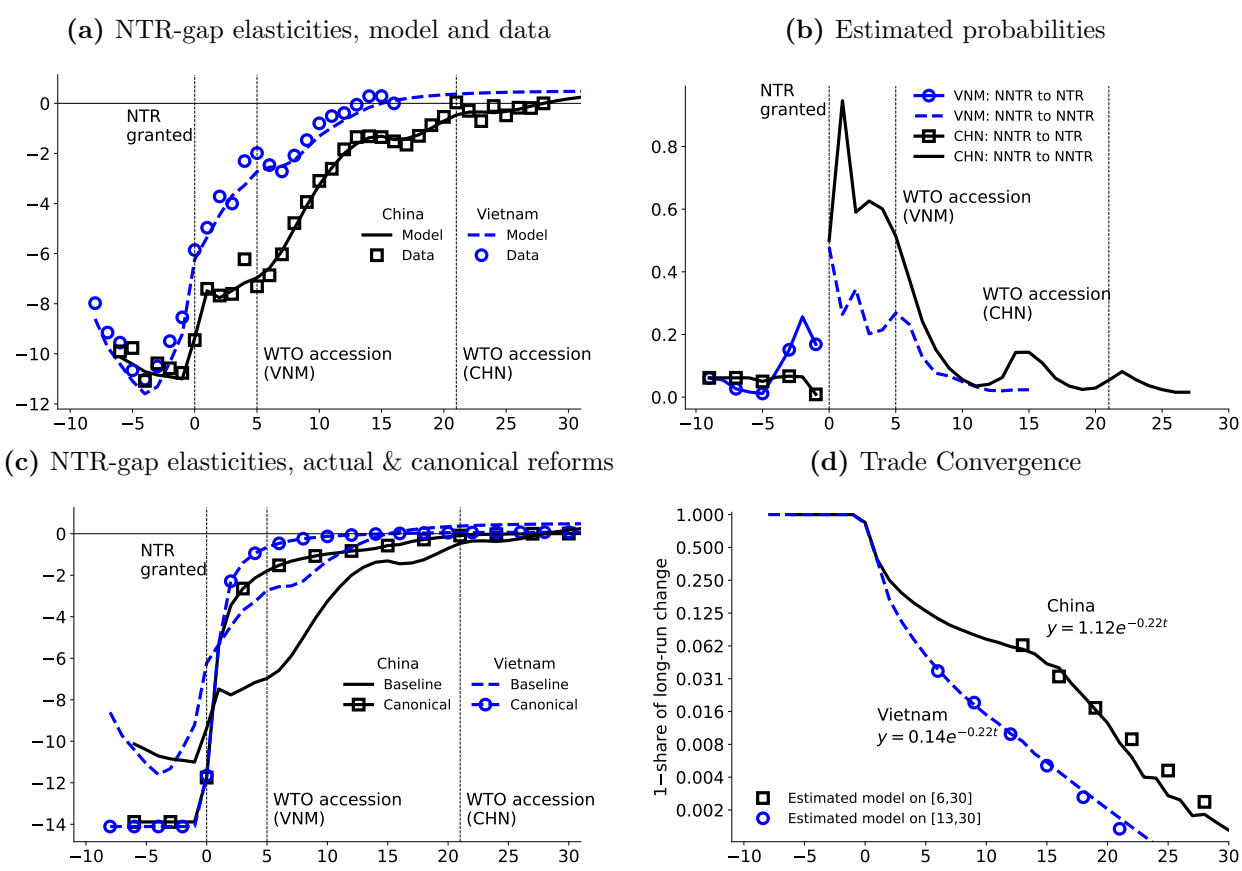
Notes: The figure plots the transition dynamics from the trade reform experiments described in section 5.3, in which the economy faces a high-persistence regime (*perm*) and a low-persistence regime (*trans*). In each regime, the tariff rate can be high (H) or low (L). In panel (e), the measured trade elasticity is defined in (4). In panels (c) and (d), outcomes are measured relative to period $t = -1$ in the canonical case $\tau_H^{\text{perm}} \to \tau_L^{\text{perm}}$.

Figure 8: Role of initial conditions: Dynamics from an embargo's end



Notes: The figure plots the transition dynamics from the experiments described in section 5.4. Panel (a): Aggregate exports after unexpectedly ending an embargo and applying either a 30-percent tariff or no tariff. Panel (b): The "gap" elasticity estimates from (20).

Figure 9: The NTR liberalizations of China (1980) and Vietnam (2002): Model



Notes: Panel (a): NTR-gap elasticities in data and baseline model. Panel (b): Estimated trade-policy transition probabilities. Panel (c) NTR-gap elasticities baseline model and the counterfactual canonical reform. Panel (d) The share of total adjustment remaining in the NTR-gap elasticities from the counterfactual canonical reform, as defined in (22).

Table 1: Trade policy regime transition frequencies (percent)

$t-1/t$		NNTR	NTR	PTA	UTPP
NNTR	jg	80.02%	18.66%	0.00%	1.31%
	j	89.37%	9.22%	0.00%	1.41%
NTR	jg	0.05%	96.94%	0.65%	2.37%
	j	0.33%	88.65%	0.18%	10.85%
PTA	jg	0.00%	8.58%	91.42%	0.00%
	j	0.00%	0.00%	100.00%	0.00%
UTPP	jg	0.01%	16.63%	0.94%	82.42%
	j	0.00%	10.35%	0.47%	89.18%5

Notes: The transition frequencies are calculated as the average of the annual transitions between and within regimes over the full sample period at the j level. Each cell in the annual transition matrix is calculated as the ratio of sum of countries that moved from regime R' to regime R between years $t-1$ and t to the total number of countries in regime R in year $t-1$, where $R = \{\text{NTR, NNTR, PTA, UTPP}\}$.

Table 2: Top five country-year transitions across regimes

From	To	jg's (# g)
NTR	NNTR	Poland-1983 (232), Poland-1984 (78), Poland-1985 (43), Afghanistan-1986 (46), Romania-1989 (119)
NTR	PTA	Canada-1989 (889), Mexico-1994 (387), South Korea - 2012 (325), Australia-2005 (241), Israel-1986 (203)
NTR	UTPP	Taiwan-1976 (280), Hong Kong-1976 (214), Israel-1976 (204), South Korea-1976 (189), Brazil-1976 (177)
NNTR	NTR	China-1980 (273), Vietnam-2002 (347), Poland-1989 (253), Soviet Union-1992 (226), Soviet Union-1993 (215)
NNTR	UTPP	Romania-1994 (32), Czechoslovakia-1992 (31), Czechoslovakia-1991 (28), Bulgaria-1992 (26), Poland-1990 (23)
PTA	NTR	Canada-1999 (205), Mexico-1999 (179), Israel-1999 (165), Australia-2009 (135)
UTPP	NTR	South Korea-1989 (403), Taiwan-1989 (400), Hong Kong-1989 (265), Malaysia-1997 (262), Portugal-1986 (214)
UTPP	NNTR	Romania-1989 (7), Romania-1990 (6), Romania-1992 (5), Romania-1993 (5), Yugoslavia-1996 (5)
UTPP	PTA	Israel-1985 (354), Mexico-1994 (342), Peru-2007 (241), Colombia-2001 (229), Dominican Republic-2007 (176)

Notes:

Table 3: Tariff levels by regimes

	Baseline Sample				Winsorized (99.99%)			
Regime	N # *jgt*'s	Mean (p.p.)	Median (p.p.)	SD (p.p.)	N # *jgt*'s	Mean (p.p.)	Median (p.p.)	SD (p.p.)
NTR	1,758,012	4.87%	2.90%	14.09%	1,757,870	4.82%	2.90%	6.66%
NNTR	19,507	38.61%	35.00%	22.56%	19,423	38.16%	35.00%	21.07%
PTA	97,583	0.78%	0.05%	2.21%	97,580	0.78%	0.05%	2.08%
UTPP	228,975	0.41%	0.00%	1.11%	228,975	0.41%	0.00%	1.11%
.	36,047	1.97%	0.12%	7.04%	36,039	1.94%	0.12%	6.78%
Total	2,104,077	4.50%	1.70%	13.58%	2,103,848	4.46%	1.70%	7.38%

Notes: .

Table 4: Tariff changes across and within regimes

		1-Year Changes				5-Year Changes			
From	To	N # *jgt*'s	Mean (p.p.)	Median (p.p.)	SD (p.p.)	N # *jgt*'s	Mean (p.p.)	Median (p.p.)	SD (p.p.)
NTR	NNTR	466	27.36	26.96	19.02	913	28.03	29.26	14.90
NTR	PTA	10,291	−3.01	−1.80	4.57	16,872	−4.21	−2.69	5.64
NTR	UTPP	29,860	−4.02	−2.90	14.53	26,911	−5.20	−3.82	6.29
NTR	NTR	1,352,360	−0.15	0.00	9.47	901,640	−0.66	0.00	15.80
NTR-Tokyo	NTR-Tokyo	180,246	−0.45	−0.17	9.66	125,516	−1.81	−0.80	34.46
NTR-Uruguay	NTR-Uruguay	197,315	−0.25	−0.04	2.46	158,828	−0.84	−0.30	2.87
NTR Non-GATT	NTR Non-GATT	974,799	−0.07	0.00	10.29	442,568	−0.24	0.00	12.18
NNTR	NNTR	10,542	−0.25	0.00	9.25	3,012	−0.44	0.00	10.02
NNTR	NTR	1,523	−27.63	−26.17	24.04	3,343	−31.48	−30.03	21.46
NNTR	UTPP	72	−29.40	−32.51	16.32	735	−36.25	−34.99	18.12
PTA	NTR	8,445	2.47	1.29	3.95	8,939	1.38	0.00	3.56
PTA	PTA	75,910	−0.12	0.00	1.34	48,483	−0.50	0.00	2.07
UTPP	NNTR	12	23.04	24.08	17.77	75	34.67	34.96	16.24
UTPP	NTR	30,244	2.90	2.09	3.81	38,291	2.34	1.59	3.42
UTPP	PTA	1,847	−0.01	0.00	1.37	7,062	−0.20	0.00	1.64
UTPP	UTPP	149,526	−0.03	0.00	1.04	82,035	−0.06	0.00	1.15
	Total	1,671,098	−0.17	0.00	8.92	1,138,311	−0.74	0.00	14.40

Notes: Tariff changes are calculated as the average over $\tau_{jgt} - \tau_{jg,t-h}$ with $h = \{1, 5\}$. For the 1-year change, GATT-Tokyo are averages for the years 1980-1987 and GATT-Uruguay 1995-2000. For the 5-year change, GATT-Tokyo are averages for the years 1984-1992 and GATT-Uruguay 1999-2004.

Table 5: Exporter-dynamics statistics and sector-level model parameters

	Target statistics				Parameters			
Country	Export part. (%)	Exit rate (%)	Incumbent prem.	Log CV exports	f_0	f_1	ξ_H	σ_z
China	28	11	2.90	2.27	0.73	0.342	3.92	1.50
Vietnam	11	15	4.44	2.91	1.57	0.657	5.89	1.69

Appendix (For online publication)

We study several features of the model and the data in this appendix. In section A we first discuss alternative reforms without uncertainty but include other forms of anticipation. In section B, we discuss the robustness of our empirical approach.

A Model

Aside from the changes between NNTR and NTR/UTPP, many tariff changes involve a phaseout. We now discuss how trade evolves with these types of tariff reforms through the lens of our model and show that short-run and long-run trade elasticities can differ quite substantially from the those from a canonical trade reform.

A.1 Anticipated reforms

To study how anticipation affects trade adjustment dynamics, we consider several variations on the canonical scenario. We consider two kinds of variations: anticipated versus unanticipated and immediate versus phaseouts. Unanticipated reforms are announced in the period in which they occur, whereas anticipated reforms are announced ten periods in advance. In immediate reforms, tariffs change by 30 p.p. in one period, and in tariff phaseout reforms, tariffs fall by 3 p.p. per year for 10 years. The linear phaseout is a key future of trade agreements, including the GATT and FTAs. In each variation, firms learn the entire path of tariffs when the reform is announced. Thus, phased-in reforms always feature some anticipation of future tariff cuts, even when the start of the reform is a surprise. None of the reforms considered in this section feature uncertainty about future tariffs. Figure A2 shows the results of these experiments. There are three main takeaways.

First, anticipation causes exports to rise before tariffs fall, increasing the short-run trade elasticity, ε_0. When the reform is anticipated, the future decrease in tariffs increases the value of exporting when the reform is announced, which causes export participation to rise. When tariffs fall, aggregate exports increase from the intensive-margin effect and the pre-reform export entry. Compared with an unanticipated reform, anticipated reforms feature more trade in $t = 0$. Thus, the measured short-run trade elasticity is higher than the short-run

canonical elasticity (Figure A2(e)).

Second, phased-in reforms cause exports to grow slower than immediate reforms, but lead to higher trade elasticities in the short and medium run. When tariffs fall gradually, the value of exporting and the export participation rate rise gradually. However, exports rise faster, relative to the rate at which tariffs change, than in immediate reforms. Export entry depends on the entire path of future tariffs, not just the current tariff rate, and firms anticipate future tariff cuts in phased-in reforms. In unanticipated phased-in reforms, the trade elasticity in period $t = 0$ is equal to the demand elasticity, θ (as in the canonical reform), but in period $t = 1$ it jumps above its long-run value. In anticipated and phased-in reforms, the trade elasticity is very high even in period $t = 0$ and exhibits similar, albeit more muted, non-monotonic behavior.

Third, anticipation does not affect the long-run level of trade but does affect the measured long-run trade elasticity. In all four reforms, export volumes converge to the same level (Figure A2(d)), but each reform has a different long-run trade elasticity. This is because the level of trade in the period before the reform begins, X_{t-1} in (4), is higher when the reform is anticipated. This is particularly true in the anticipated, immediate reform, where more than three-quarters of the long-run growth in trade occurs before tariffs change at all, which results in a much lower long-run trade elasticity than in the other versions of the experiment. Of course, if we measure the change in trade relative to the announcement date, as in Figure A2(g), the long-run elasticity is the same across experiments. Empirical approaches that attributes this anticipatory growth to pre-trends will generate even smaller trade elasticities.

B Data

B.1 Regime classification

To classify each triplet jgt into one of the four regimes, we incorporate information from the rate of provision and country subcode aggregation level of the census trade data. In particular, we classify jgt's into each of the regimes when at least 50 percent of the import value of the triplet is attributed to that regime. We define regimes at the jgt-rate of provision-

country subcode level as follows:

- *NTR*: Rates of provision 10, 61 or country subcodes C, L, and K.
- *NNTR*: Rates of provision 62.
- *PTA*: Country subcodes specified in Table A3 or rates of provision 18, 19, or 64 for country-years in Table A3.
- *UTPP*: Country subcodes specified in Table A1 or rates of provision 18, 19, or 64 for country years with at least one observation under country subcode A, A*, A+ or 4 (GSP).

This procedure allows us to classify 98.4 percent of the jgt triplets (2,587,514/2,629,684). Panel a of Figure A6 plots the import share and distribution of applied tariffs for the unclassified jgt's. It indicates that these represent a small fraction of imports and are mostly characterized by low tariffs until 2017, when the U.S. began its protectionist turn. The Trump's tariffs on imports of steel and aluminum, and, especially, on imports on China, was mostly carried out under provision rates 69 and 79, which are not used by any of our 4 regimes. This justifies why we focus our sample period until 2017.

The imports of most of the jgt's that are classified according to one of our four regimes are completely characterized by their regime. This can be seen in Panel b of Figure A6, which plots the import share of the attributed jgt regime. The import share of the regime is always very close to 1, especially in the case of the NTR and NNTR regime. The fact that the regime of some jgt's does not account for all imports is mostly due to product aggregation. Consider for example electric motors incl. universal A.C./D.C. motors (sitc code 71621) imported from Canada in 1990. We classify this triplet as "PTA". However, the "PTA" import share of this jgt is only 65 percent, with the "NTR" share being 34 percent and 1 percent unclassified. Table shines some light on this: There are XX HTS-8 products associated with this SITC code, of which some provide lower than NTR rates under the U.S.-Canada FTA while others already provide zero tariffs under the NTR regime.

B.2 Estimation methods

We estimated trade elasticities using a local projection approach in section 3. We now show two additional common methods to estimate the dynamic pattern of trade adjustment yield quite similar results. Our first approach, which we call the the h-on-h method, measures the ratio of the cumulative change in trade and tariffs. Our second approach is a partial adjustment model known as the error-correction model. Partial adjustment models have been used widely to account for trade dynamics.

The h-on-h method. First, we consider a direct application of (4) to the data, regressing changes in trade after h periods on changes in tariffs after h periods while using fixed effects to control for confounding covariates. The estimating equation, which we call the *h-on-h* specification, is:

$$\Delta_h v_{jgt} = -\beta_h^{hh} \Delta_h \tau_{jgt} + \delta_{jt} + \delta_{gt} + u_{jgt}. \tag{23}$$

where $\Delta_h x_t = x_{t+h} - x_{t-1}$ and β_h^{hh} is the estimator of the h-horizon trade elasticity under the h-on-h approach. We provide estimates for horizons $h = [0, 14]$ for China and Vietnam around the NTR granting and more generally for the whole dataset. Standard errors are clustered at the jg level. Figure A1 shows that the pattern of adjustment for China and Vietnam are quite similar with this approach as our h-on-1 approach. For other transitions, the h-on-h shows more gradual growth than the h-on-1 approach.

The error correction method. Our third specification, which we call *ECM*, is a version of the error correction model (Hendry et al., 1984; Hooper et al., 2000; Gallaway et al., 2003; Alessandria and Choi, 2021). This approach allows us to recover a parametric estimate of the long-run trade elasticity by specifying trade flows as an autoregressive process of order one and allowing for lagged effects of tariffs. We estimate:

$$\Delta_0 v_{jgt} = \beta_0^{ecm} \Delta_0 \tau_{jgt} + \gamma(v_{jg,t-1} - \beta_\infty^{ecm} \tau_{jg,t-1}) + \delta_{gt} + \delta_{jt} + \delta_{jg} + u_{jgt}, \tag{24}$$

where the dependent variable is the one-year log difference in import values between t and $t-1$, and the right-hand side includes the one-year change in tariffs, lagged tariffs, lagged log imports and a set of fixed effects. The one-year, or short-run, elasticity, β_0^{ecm}, is the

coefficient on the contemporaneous tariff change. The long-run elasticity, β_{∞}^{ecm}, is pinned down by the ratio of the coefficient on lagged tariffs to the coefficient on lagged imports or speed of adjustment. The fixed effects are the same as in (5) except for the inclusion of variety fixed effects, δ_{jg}, which capture each variety's average import level. This is warranted because, in contrast with the previous two specifications, jg fixed effects do not cancel out with the use of changes on the left hand side since once can rewrite the ECM specification in levels.[23] One important difference between the two local projection specification, 'h on 1" and h-on-h, is that the former can fix the reform period at t, while the "ECM" approach averages over all changes.

B.3 Data construction

We now describe how we classify good regimes in our data

B.4 Alternate Samples

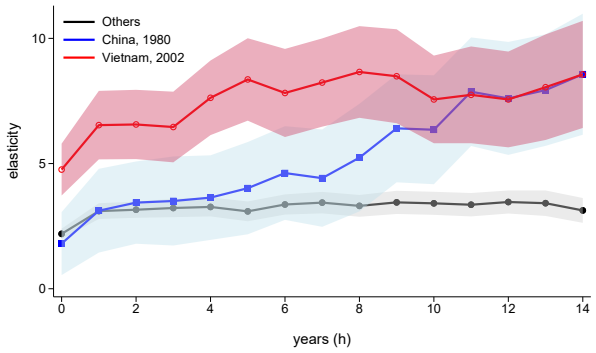
We now describe the properties of tariffs and trade for alternative samples. Specifically, we report moments for i) all data without conditioning and ii) An aggregation that takes into account the heterogeneity across products.

C Additional figures

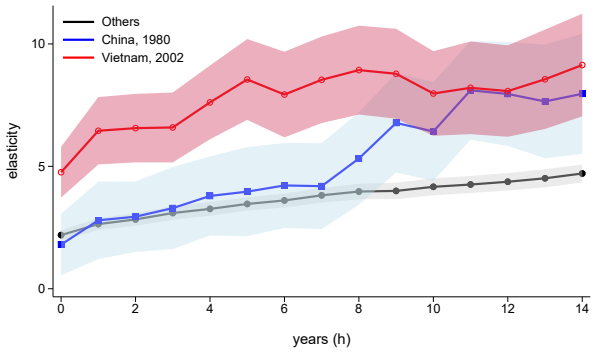
[23] This is, (24) can be written as $v_{jgt} = \beta_0^{ecm}\tau_{jgt} + (1+\gamma)v_{jg,t-1} - (1+\gamma\beta_{\infty}^{ecm})\tau_{jg,t-1} + \delta_{gt} + \delta_{jt} + \delta_{jg} + u_{jgt}$.

Figure A1: Conventional estimates

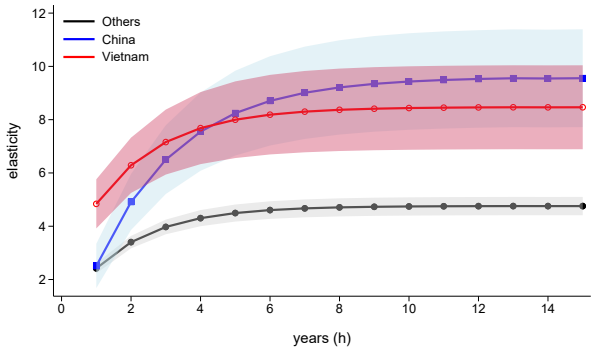
(a) Trade elasticities, h-on-1



(b) Trade elasticities, h-on-h



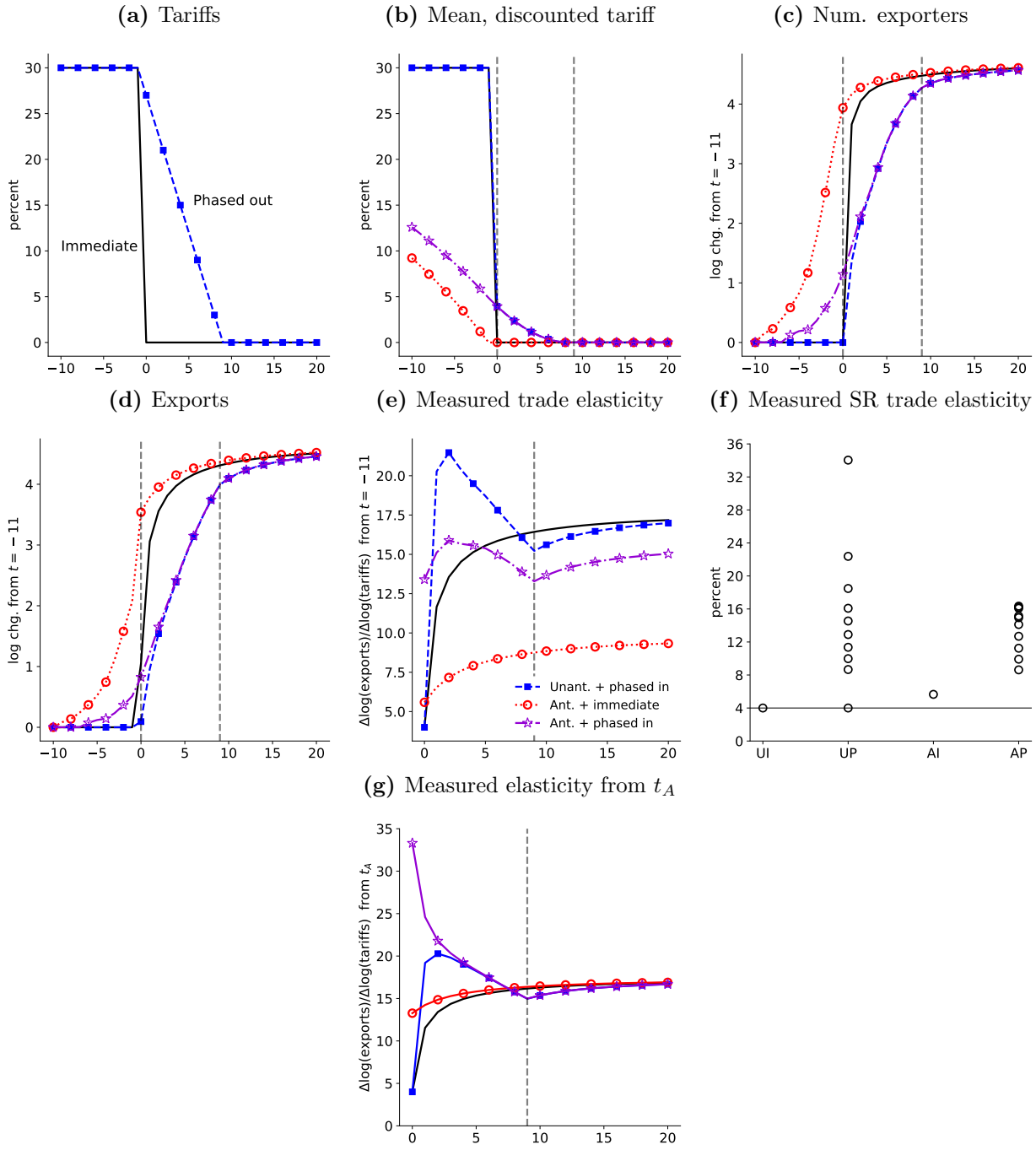
(d) Trade elasticities, ECM



(c) UPDATE: Trade elasticities, h-on-1 regimes & countries

Notes: Estimates interact the coefficients of interest of the h-on-h and the h-on-1 specifications with a categorical variable for China, Vietnam, and others. Estimates in panel (c) interact the coefficients of interest of the ECM specification with a categorical variable for China, Vietnam, and others. The standard errors that construct the 95 percent confidence intervals are clustered at the jg level.

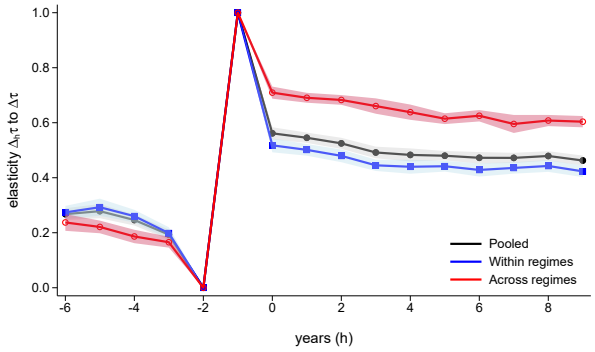
Figure A2: Model responses to anticipated vs. unanticipated reforms



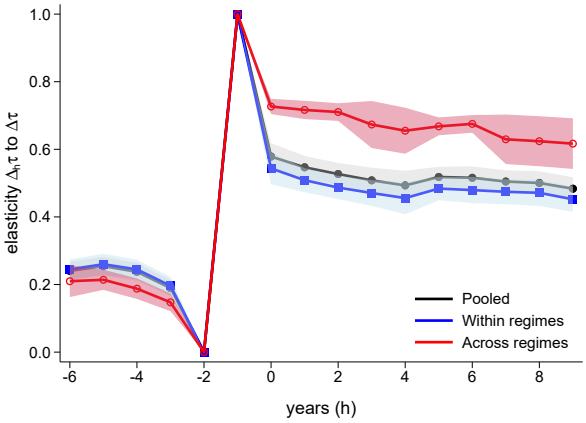
Notes: Figure shows transition dynamics in experiments described in section A.1. Panel (a): observed path of tariffs in immediate and phased-in reforms. Panel (b): Mean discounted expected tariff in unanticipated/immediate, anticipated/immediate, unanticipated/phased-in, and anticipated/phased-in reforms. Panel (c): number of exporters. Panel (d): aggregate exports. Panel (e): measured trade elasticity defined in (4). In panels (c) and (d), outcomes measured relative to period $t = -11$ to illustrate how they begin to respond after reform is announced, but before tariffs actually change.

Figure A3: Tariff autocorrelation and fixed effect structure

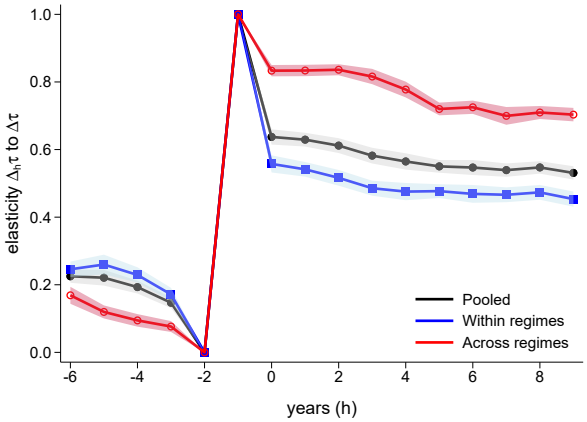
(a) jt, gt FEs



(b) jt FEs



(c) gt FEs



(d) No FEs

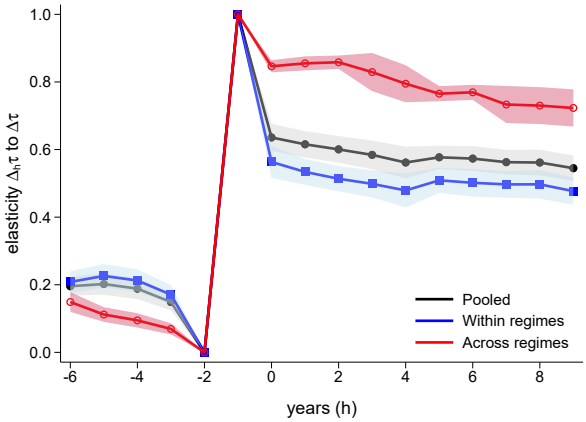
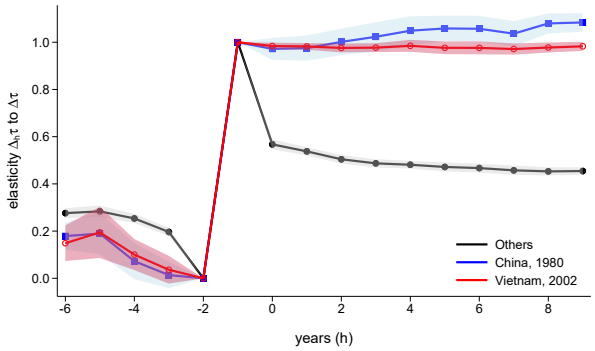
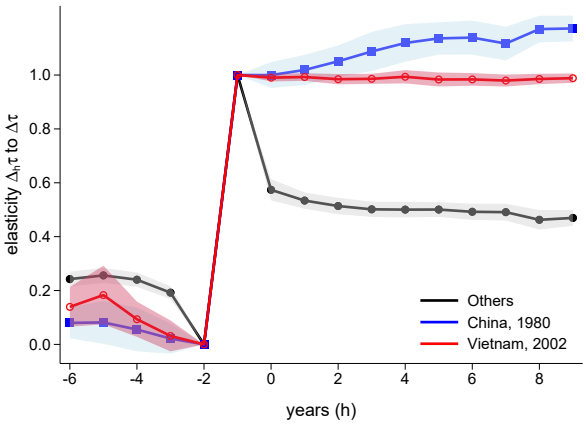


Figure A4: Tariff autocorrelation China and Vietnam and Fixed effects

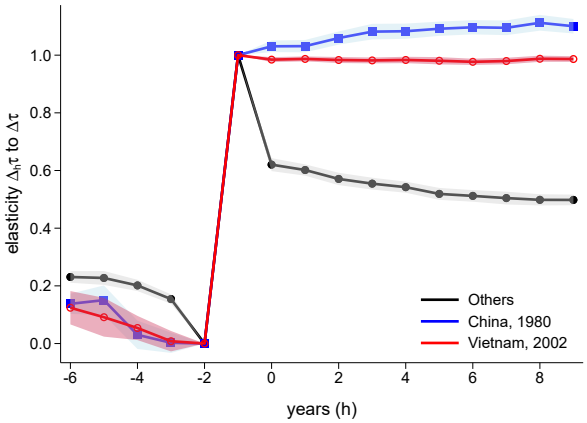
(a) jt, gt FEs



(b) jt FEs



(c) gt FEs



(d) No FEs

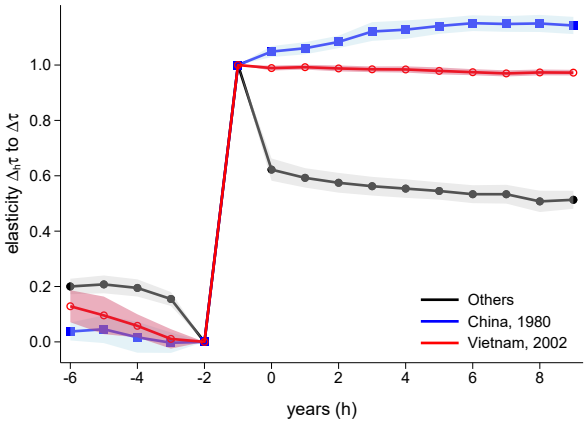
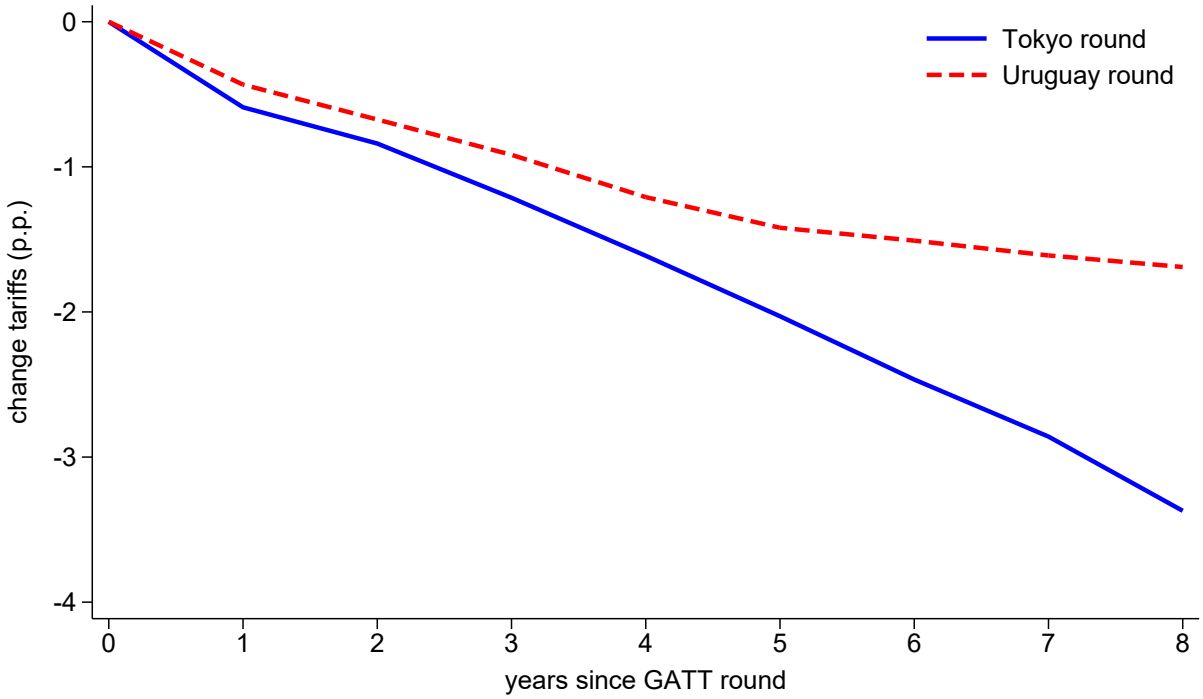


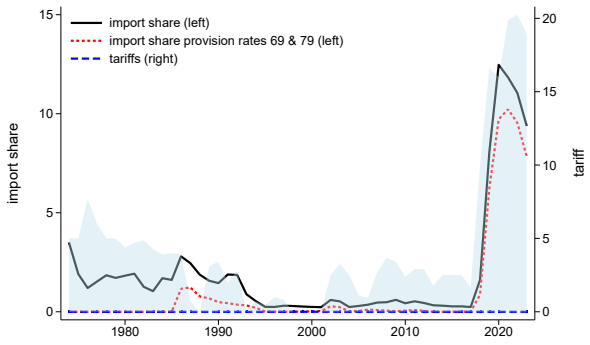
Figure A5: Tariff Phaseouts from GATT rounds



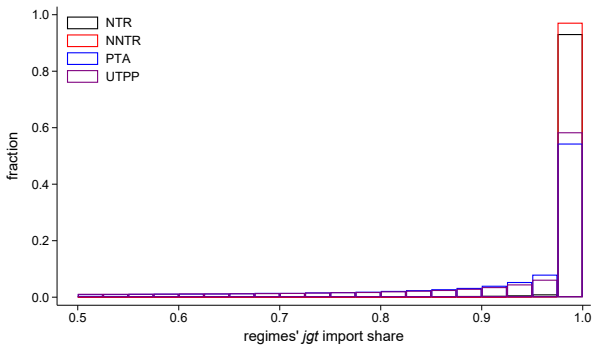
Notes: The figure shows the average tariff cut across varieties for each year, relative to the period before the Tokyo (1979) and Uruguay (1994) GATT rounds.

Figure A6: Regime Classification

(a) Unclassified *jgt* observations



(b) *jgt* Regime Import Share



D Additional tables

Table A1: List of Unilateral Trade Preference Programs

Year	Program	Code
1976	Generalized System of Preferences (GSP)	A, A*, A+, 4
1982	Caribbean Basin Economic Recovery Act (CBERA)	E, E*, 9
1992–2014	Andean Trade Preferences Act (ATPA)	J, J*, J+
2001	African Growth and Opportunity Act (AGOA)	D
2004	Caribbean Basin Trade Partnership Act Initiative (CBTPA)	R
1999	Duty-free Code for West Bank and Gaza Strip	N
2017	Nepal Preference Program	NP

Notes: The *Year* column is the first (and last) year in which we observe the corresponding country sub-code in the U.S. Census trade data, reported in the *Code* column.

Table A2: List of countries in a PTA with the United States

Year	Country	Code	Year	Country	Code
1985	Israel	IL	2006	Singapore	SG
1988	Canada	X, CA	2007	Peru	PE
1994	Mexico	MX	2010	Costa Rica	P, P+
2001	Colombia	CO	2010	El Salvador	P, P+
2001	Jordan	JO	2010	Dominican Republic	P, P+
2004	Australia	AU	2010	Guatemala	P, P+
2004	Chile	CL	2010	Honduras	P, P+
2006	Bahrain	BH	2010	Nicaragua	P, P+
2006	Morocco	MA	2010	South Korea	KR
2006	Oman	OM	2012	Panama	PA

Notes: The *Year* columns report the first full year the agreement went into effect. The *Code* columns are the country sub-codes reported in the U.S. Census trade data.

Table A3: Example of the Regime Classification

Panel a: Census Data for SITC good

Provision Rate	Country Subcode	Tariff (%)	Import Share (%)	Regime
10	0	0.00	5	NTR
13	0	0.00	1	.
16	0	0.00	0	.
18	B	0.00	10	PTA
18	C	0.00	0	PTA
61	0	4.12	28	NTR
64	0	3.87	23	PTA
64	X	3.00	31	PTA
79	0	0.00	0	.
79	X	0.00	1	PTA

Panel b: 1990 HTS Tariff Schedule for corresponding HTS8 Products

HTS-8 Product	NTR rate (%)	PTA rate (%)
85011020	10	8
85011040	6.6	5.2
85011060	4.2	3.3
85012020	4.2	2.5
85012040	5	3
85012050	5	3
85012060	3.7	2.2
85014020	4.2	3.3
85014040	5	4
85014050	5	4
85014060	3.7	2.9
85015120	4.2	3.3
85015140	5	4
85015150	5	4
85015160	3.7	2.9
85015240	3.7	2.9
85015280	Free	
85015340	Free	
85015360	4.2	2.5
85015380	4.2	2.5

Notes: