The GATT vs Inflation: Tokyo Drift*

Andrew Greenland†
Elon University

James Lake‡
University of Tennessee

John Lopresti§
William & Mary

June 9, 2023

Preliminary and Incomplete
Latest Version

Abstract

Despite the negotiated tariff liberalization in the 1979 Tokyo GATT Round, the US aggregate tariff rate drifted upwards over the following decade. In stark contrast, the aggregate tariff rate fell sharply in the years prior to the Tokyo Round despite no legislated changes in US tariff policy. We shed light on these puzzling dynamics by reconstructing the US tariff code annually between 1972 and 1988 and document a two-phase liberalization. From 1972 to 1979, inflation eroded the ad valorem equivalent rate of specific tariffs – which account for up to 35% of tariff protection in our sample – resulting in an "accidental" liberalization. Between 1980 and 1988, a change in the composition of US imports masked a four percentage point reduction in legislated tariffs resulting from Tokyo Round GATT negotiations. To emphasize the aggregate importance of specific tariffs in this era, we embed them in a simple CES framework and extend existing hat algebra techniques to incorporate specific tariffs. Our counterfactual analysis shows the inflationary erosion of specific tariffs yielded greater welfare gains than the GATT-mandated phaseout of legislated tariffs.

*We thank Treb Allen, Costas Arkolakis, Rafael Dix-Carneiro, Lydia Cox, Teresa Fort, Matt Grant, Doug Irwin, Chris Meissner, Emanuel Ornelas, Ralph Ossa, Nina Pavcnik, Martin Rotemberg, Georg Schaur, Bob Staiger, Claudia Steinwender, Felix Tintelnot, Sharon Traiberman, David Weinstein, and Gerald Willman for helpful discussions on this project, as well as seminar and conference participants at Dartmouth College, Duke University, North Carolina State University, the 2023 CESifo Global Economy Conference, the Columbia University & NYU international trade workshop, the Fall 2022 Midwest Trade Meetings, and the Workshop on Trade Policy and Trade Agreements at Bielefeld University.

†E-mail: agreenland@elon.edu
‡E-mail: jlake7@utk.edu
§E-mail: jwlopresti@wm.edu

1 Introduction

In exploring the effects of trade on US workers, firms, and communities, researchers have overwhelmingly focused on the past 30 years. This is perhaps natural, given the rapid growth of trade in the era of "hyper-globalization".[1] Tariff liberalization played a central role in this era, through the completion of the Uruguay Round of negotiations under the General Agreement on Tariffs and Trade (GATT), the proliferation of regional trade agreements, and the accession of economic powers China and Russia to the World Trade Organization.

From an empirical standpoint, however, this focus comes with an obvious caveat: by 1990, much of the post-war liberalization project in the US was already complete. Over the past century, the US ad valorem equivalent (AVE) tariff level, defined as the ratio of duties collected to total imports, peaked at approximately 20% in the years following the Tariff Act of 1930. By 1990, this had fallen to 3.3%.[2] To the extent that trade has shaped the evolution of the US economy, it likely did so in the decades prior to the ones that scholars have chosen to emphasize.

A primary reason for the relative paucity of work on prior eras is the lack of product-level import tariff data, which only become available with the advent of the Harmonized System (HS) in 1989. Indeed, Anderson and Van Wincoop (2004, p.693) argue that "the grossly incomplete and inaccurate information on policy barriers available to researchers is a scandal and a puzzle". In this paper, we take a step towards addressing this shortcoming by developing an algorithm to reconstruct the US tariff code between 1972 and 1988, a period during which the aggregate AVE level fell by approximately 25% *more* that in the years between 1990 and 2016. Our procedure relies only on publicly available data used extensively by trade economists, covers the vast majority of US imports during this period, and yields highly accurate estimates of annual legislated tariff rates. Strikingly, we demonstrate that our algorithm yields nearly identical rates to those which would be obtained with a perfect digitization of PDF tariff schedules. A primary contribution of this paper is to make these data available to other researchers.

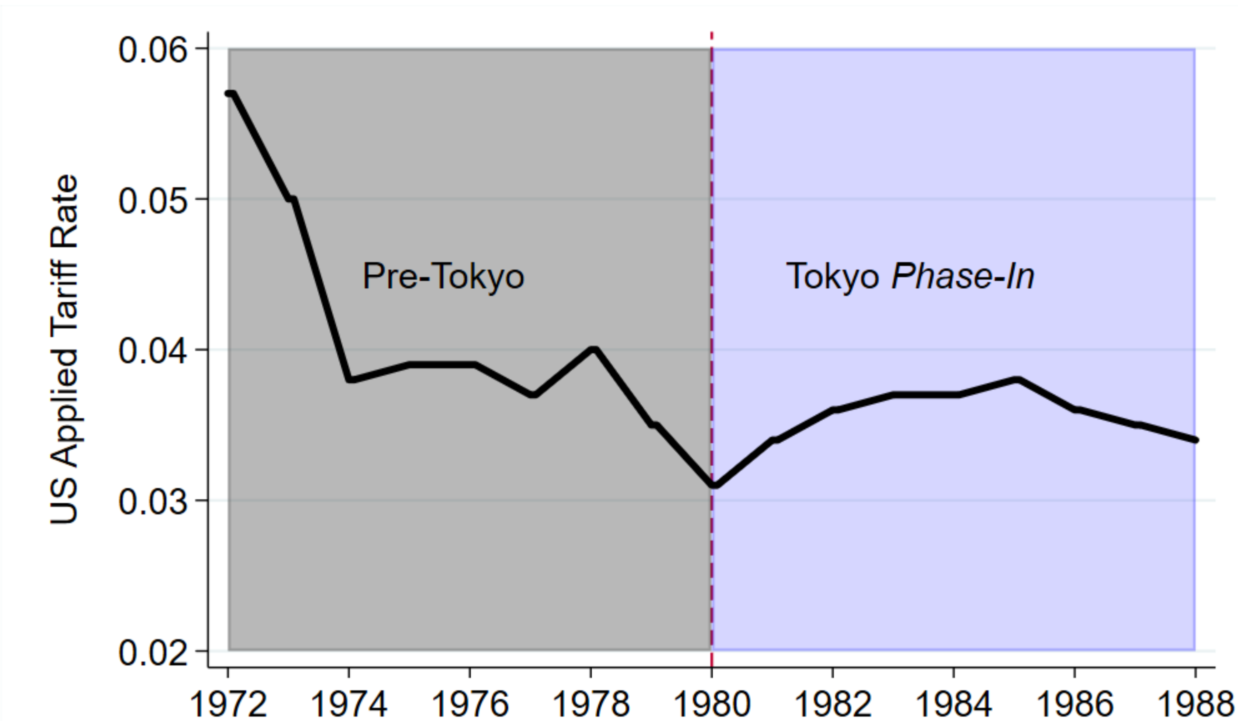
A second contribution is to use the data to explain a puzzle in the evolution of US tariffs in this era. The Tokyo Round of the GATT was completed in 1979, with negotiated tariff reductions phased in over the following eight years. This represented the first major change

[1] In the words of The Economist, "The golden age of globalization, 1990-2010, was something to behold": https://www.economist.com/leaders/2019/01/24/the-steam-has-gone-out-of-globalisation.

[2] https://www.usitc.gov/documents/dataweb/ave_table_1891_2016.pdf

in the US tariff code since the tariff reductions mandated by the 1967 Kennedy Round of the GATT were completed in 1972. One would thus expect relative stability in average US tariffs throughout the 1970s and a subsequent decline throughout the 1980s. However, Figure 1 depicts the US aggregate AVE tariff as reported by the US International Trade Commission (USITC) and shows that the *opposite* is true.[3] That is, the average tariff level fell sharply over the decade despite the absence of meaningful changes to trade policy for most of the 1970s, and despite the successful completion of the Tokyo Round negotiations in 1979, average tariffs *rose* slightly in its wake. Whether trade policy matters for economic outcomes is, of course, a question of enormous interest to researchers (Goldberg and Pavcnik, 2016; Caliendo and Parro, 2021). To this end, Figure 1 raises a fundamental question: how does trade policy map into observed trade barriers as typically measured by trade economists?

Figure 1: Aggregate US Ad Valorem Equivalent Tariff Rate



Notes: US Ad Valorem Equivalent (AVE) Tariff Rate defined as duties collected divided by total imports. Data from USITC.

Given the magnitude of the liberalization during the 1972-1988 period and the relative lack of work studying it, answers to the puzzles posed by Figure 1 are important in their

[3]Accessed from the USITC website.

own right.[4] However, these data are also valuable as a foundation for future research into an era of considerable economic importance. An enormous literature is dedicated to exploring the relationship between US manufacturing decline and international trade, again focusing primarily on the years after 1990 (Autor et al., 2013; Pierce and Schott, 2016). However, this largely ignores the fact that US manufacturing employment peaked in the late 1970s and has declined continuously in the decades since. Thus, a full accounting of the role of trade in manufacturing's decline must begin decades prior to the majority of existing work. This is also an era in which the college wage premium, women's labor force participation, and union membership rates all change dramatically. The relationship between these phenomena and trade is of first order importance. We hope these data can be used by researchers to explore such relationships more fully.

Using our estimates of legislated tariffs, we show that the aggregate pattern in Figure 1 is the product of two distinct changes. First, between 1972 and 1979 there is an "accidental liberalization" caused by the inflationary erosion of protection afforded by specific (per-unit) tariffs. Ad valorem equivalent rates on goods facing specific tariffs fall from 7% to 3% in these years, driving the aggregate decline shown in Figure 1.[5] Second, the Tokyo Round negotiated tariff reductions are phased in between 1980 and 1988 as a step-wise decline in annual tariff rates, largely according to the "Swiss Formula".[6] This product-level liberalization is masked, however, by the changing composition of imports among incumbent products. Holding import shares fixed among continuing products, we find that tariffs are 27% lower in 1979 than in 1972 and fall another 25% with the tariff phase-ins scheduled under the Tokyo Round. While nearly every sector experiences tariff cuts, raw materials and agriculture are primarily impacted in the 1970s while remaining sectors are primarily affected by Tokyo Round phaseouts throughout the 1980s.

As a final contribution, we extending existing hat algebra and exact hat algebra tech-

[4] Much of the work on the Tokyo Round was conducted during the early years of its phase in, with a focus on general equilibrium estimates of the agreement's effects on the structure of production across member countries (Brown and Whalley, 1980; Deardorff and Stern, 1981, 1983) or on the nature of the negotiations themselves (Ahmad, 1978; Chan, 1985).

[5] Goods facing specific tariffs account for approximately one-third of all imported products prior to 1979. The importance of specific tariffs for tariff levels in this era was noted in more aggregate data by Van Cott and Wipf (1983).

[6] The Tokyo Round was the first use of Swiss Formulas for tariff reduction. The formula defines a final tariff Z as $Z = \frac{AX}{A+X}$, where X is the initial tariff and A is a coefficient representing the maximum final tariff rate across all products. In the Tokyo Round negotiations, A was set at 0.14, though the formula was not followed for all products. For more details, see: https://www.wto.org/english/tratop_e/agric_e/agnegs_swissformula_e.htm.

niques to account for specific tariffs in a standard CES framework and use these techniques to perform counterfactual analyses. Given the dependence of the counterfactual analyses on estimates of import demand and export supply elasticties, we generalize the Feenstra (1994) and Broda and Weinstein (2006) elasticity estimator to account for specific tariffs and show how the traditional estimator leads to biased elasticity estimates. Further underscoring the quantitative importance of specific tariffs, we then show that while both the inflationary erosion of specific tariffs and legislated tariff cuts cause quantitatively important declines in US import prices, inflation accounts for the majority of these effects. That is, the "accidental liberalization" is a more important determinant of the welfare consequences of trade in this era than the negotiated tariff cuts in the Tokyo GATT Round.[7] To the best of our knowledge we are the first to document the importance of specific tariffs relative to ad valorem tariffs.

Our work is related to three distinct strands of literature. First, we add to the numerous studies focused on quantifying the effects of trade policy on US domestic economic outcomes. Studies of the 1989 Canada-US Free Trade Agreement (Trefler, 1993; Kovak and Morrow, 2022), NAFTA (Caliendo and Parro, 2015; Hakobyan and McLaren, 2016), the US granting PNTR to China in 2001 (Pierce and Schott, 2016; Handley and Limao, 2017; Greenland et al., 2019), the 2001 Bush steel tariffs (Cox, 2022; Lake and Liu, 2021), and the Trump-era tariffs (Fajgelbaum et al., 2020; Flaaen and Pierce, 2021) all rely on discrete changes in disaggregate tariff data to establish causal links between trade policy and economic outcomes. By contrast, we emphasize the importance of both the inflationary erosion of specific tariffs and legislated tariff changes as sources of liberalization and extend existing hat algebra and exact hat algebra techniques to incorporate specific tariffs. Further, in hopes of facilitating subsequent research in this era, we also construct the first publicly available dataset of TSUSA legislated tariffs at the annual level, designed to complement the US import data constructed by Feenstra (1996).[8] Given the considerable growth in imports, substantial declines in tariffs, and labor market turmoil during this time period, we believe these data and techniques will prove valuable for other researchers.

Second, our study contributes to the literature focused on the impacts of additive trade barriers. Hummels and Skiba (2004) and Feenstra and Romalis (2014) emphasize the importance of per-shipment trade costs in determining both trade flows and export quality.

[7]Beyond underscording the importance of specific tariffs as a driver of trade liberalization in this era, the model also yields tractable predictions about both the volatility of import prices and the passthrough of foreign price shocks into domestic prices in the presence of per-unit trade costs. In ongoing work, we are quantify the role of specific tariffs in these channels.

[8]These data can be accessed via The Center for International Trade Data, housed at UC Davis.

Irarrazabal et al. (2015) detail a procedure to estimate them, while Sørensen (2014) shows that the gains from trade are larger in the presence of these types of trade barriers. Still other studies emphasize the importance of per-unit trade costs in shaping tariff levels and import patterns during the first half of the 20th century US (Crucini, 1994; Irwin, 1998; Bond et al., 2013; Greenland and Lopresti, 2022). In this paper, we show that such trade costs are quantitatively important for trade flows far more recently than has been previously demonstrated. Specific tariffs cover more than one-third of products in 1972, and more than 25% of goods on the eve of the transition to the HS system.[9]

Third, we contribute to the literature focused on estimating import demand elasticities. While recent work by Boehm et al. (2020) employs an IV strategy, much of this literature has taken a structural approach. Studies stemming from Caliendo and Parro (2015) require multiple countries' bilateral trade data to recover import demand elasticities. However, no such data are available during our sample. By contrast, Feenstra (1994) presents an estimation strategy suitable to obtaining structural estimates of demand and supply elasticities in the absence of independent instruments using import data from a single country. Subsequent studies have detailed computational improvements based on theoretical insights regarding feasible parameters Broda and Weinstein (2006) and alternative estimation techniques Soderbery (2015). Here, we build on these studies by generalizing the Feenstra (1994) estimator to account for specific tariff changes and detail biases caused by their omission. Researchers should be aware of such biases when estimating elasticities, especially in settings in which goods are reliant on specific tariffs for protection.

The paper proceeds as follows. Section 2 describes the data and algorithm used to reconstruct the TSUSA tariffs from 1972-1988 and evaluates the accuracy of the estimates. Section 3 provides an explanation for the puzzle surrounding the disconnect between the aggregate US AVE tariff rate and conventional wisdom about US tariff policy between the Kennedy and Tokyo Rounds. Section 4 embeds specific tariffs in a CES framework, extends existing hat algebra techniques to incorporate specific tariffs, estimates elasticities, and presents counterfactual welfare analyses. Section 5 concludes.

[9] As of 2020, specific tariffs are still a highly important form of protection in US agriculture. Moreover, some countries remain dependent on specific tariffs for protection across the entire tariff schedule. Over 85% of Switzerland's tariff code, for example, consists of specific tariffs.

2 Data

Between 1972 and 1988, the global trading system operated under two GATT regimes. Beginning in 1968, the Kennedy Round of GATT negotiations was enacted and tariffs were phased in over four years, reaching their negotiated final levels in 1972. Upon completion of the Tokyo Round in 1979, negotiated tariff cuts were implemented with a phase-in schedule that lasted until 1987. Legislated tariff schedules for the US in this period, including phase-ins, were covered by the TSUSA.

Currently, no dataset contains annual product-level US legislated tariffs for these years. As such, work relying on tariff levels in this era has taken an indirect approach: US imports and exports are recorded annually by the US Census in the Annual Import Data Bank (IDB) and have been detailed at length by Feenstra (1996). Product-level ad valorem equivalent rates can thus be calculated by dividing duties collected by import values. However, such an approach ignores the distinction between tariffs specified in ad valorem terms and those specified in specific (per-unit) terms. One resulting problem is that researchers cannot determine whether a falling AVE rate is driven by a reduction in the legislated tariff rate or by price increases in the presence of specific tariffs. As price levels doubled throughout the 1970s and specific tariffs were pervasive, this is a non-trivial concern. To address this shortcoming, we construct a product-level dataset of US tariff rates using a novel estimation procedure that captures both the ad valorem and specific components specified by legislation.

An obvious alternative to our approach would be to digitize the annual US tariff schedules directly from available PDFs. However, this approach has an important drawback: in many years the IDB does not contain the unit of quantity (e.g. pounds, tonnes etc.). In these cases, digitizing the ad valorem equivalent of a specific tariff would require an assumption about the unit of quantity in the IDB because one needs to know the unit value *in the unit of quantity* specified in the TSUSA. In contrast, our estimation approach does not suffer from this problem because it does not use any data from the TSUSA and remains agnostic about the unit of quantity in the import data.[10]

[10] Additionally, digitizing the TSUSA schedules is a time-intensive task that is subject to at least some degree of human error. Each TSUSA schedule contains roughly 700 pages of tariffs covering up to 7000 products annually.

2.1 Estimating Tariffs 1972-1988

The import data we use to estimate tariffs are taken from the IDB.[11] These data include total duties collected as well as import values, dutiable values, and (up to) two measures of quantity. The data are highly disaggregate: each observation represents exports by an exporting country of a TSUSA seven-digit product to a US port in a year t subject to tariffs defined by "rate provision code" r. Since tariffs are applied at the five-digit good level g, the rate provision code r applies to all seven-digit imports within the five-digit good and indicates whether the duty is ad valorem, specific, or compound – i.e., a combination of ad valorem and specific duties.

Table 1 describes the two-digit rate provision codes in the data. A first digit of values 1-4 describes whether the tariff is duty free, specific, ad valorem, or compound. For non-duty-free products, a second digit of 1 or 2 indicates that the tariff is the MFN normal trade relations (NTR) tariff from column 1 of the US tariff schedule or the non-NTR tariff from column 2, respectively. The vast majority of imports fall under one of the seven potential combinations of these rate provision codes. For example, rate provision code 22 indicates a non-NTR specific tariff, while rate provision code 31 indicates an NTR ad valorem tariff. The codes 5*, 7*, *3, *4 and *8 capture various infrequently used types of tariffs.

Table 1: Rate provision codes (r)

First Digit	Duty Type	Second Digit	Rate
1*	Duty Free	*1	Col 1
2*	Specific Duty	*2	Col 2
3*	Ad Valorem	*3	Col. 1 Exceptions
4*	Compound	*4	Col. 2 Exceptions
5*	Minimum	*8	Cuba Special
7*	Special		

In general, the total duties collected on variety v (defined in the data as an exporter by seven-digit product by port triple) of good g under rate provision r at time t are

$$TD_{vgrt} = \tau_{grt} DV_{vgrt} + f_{grt} Q_{vgrt}. \tag{1}$$

Here, τ_{grt} and f_{grt} are, respectively, the unknown ad valorem and specific tariffs, while DV_{vgrt}

[11] Feenstra (1996) reports these data at the country-product-year level by aggregating over (i) dutiable and non-dutiable imports and (ii) ports. Thus, we use the raw ASCII files to access the most disaggregate data.

and Q_{vgrt} represent dutiable import value and the quantity of imports. We again note that the data in the IDB are reported at the variety-level while legislated tariffs are specified at the more aggegate five-digit level g.

Absent measurement error, one would need at most two non-collinear variety-level observations of a good in a given year to calculate its tariff(s).[12] For five-digit product g facing rate provision code r in year t, one would calculate

$$f_{grt} = \frac{TD_{vgrt}}{Q_{vgrt}} \text{ for } r = 2^* \tag{2}$$

$$\tau_{grt} = \frac{TD_{vgrt}}{DV_{vgrt}} \text{ for } r = 3^* \tag{3}$$

$$TD_{vgrt} = \tau_{grt} DV_{vgrt} + f_{grt} Q_{vgrt} \text{ for } r = 4^*. \tag{4}$$

For ad valorem and specific tariffs, only a single variety-level observation would be needed using (2) and (3). But two variety-level observations would be needed to calculate τ_{grt} and f_{grt} for compound tariffs since (4) is a single equation with two unknowns τ_{grt} and f_{grt}.[13]

However, multiple forms of measurement error make direct calculation infeasible. First, the IDB data round import values and quantities to the nearest integer. Second, database entry error may occur – occasionally, for instance, the units of quantity may be recorded inconsistently within a rate code. Third, countries may be moved from Column 2 to Column 1 status within a year and hence face different tariffs during the year.[14] This creates problems because the IDB only records imports annually.

An additional complication is that many products with specific tariffs report two distinct quantity values for a given observation, each corresponding to a different unit of measurement. Without information on which unit is used in assigning duties, there are thus two possible values for the specific tariff.[15] In such cases, one needs two varieties to calculate the correct f_{grt}. One would then calculate a value of f_{grt} using (3) for each of the two units of

[12] This standard is easily met in the data – it fails for a mere 0.2% of observations covering 0.1% of imports by value.

[13] One would obtain identical τ_{grt} and f_{grt} for any chosen variety in equations (2)-(3) and any two varities in equation (4). Additionally, the two versions of equation (4) defined by the two varities must be linearly independent – that is, they must define two non-parallel lines in τ_{vgt}-f_{vgt} space.

[14] One such is example is China gaining NTR status with the US in March of 1980.

[15] For example, TSUSA product 37315 covers non-ornamented, wool neckties for men and boys. In 1978, the rate for this was 10.5% ad valorem plus $0.375 per pound. The IDB data report two different quantities for each observation of this product. According to the TSUSA schedule, the quantities reflect dozens of ties as well as pounds of imports. However, the IDB does not indicate which value of quantity measures dozens of ties and which measures pounds of ties.

quantity and identify the correct f_{grt} as the one that is constant across the two varieties.[16] Similarly, one would need two sets of variety pairs to calculate the correct τ_{grt} and f_{grt} for products with compound tariffs. In this case, one would separately solve the simultaneous equations given by (4) for each unit of quantity. Only one unit of quantity would yield a constant τ_{grt} and f_{grt} across the two sets of variety pairs. These would be the correct τ_{grt} and f_{grt}.

Ultimately, we can only estimate the five-digit MFN tariffs τ_{grt} and f_{grt} with noise. However, Section 2.2 will show the resulting estimation error is negligible for the vast majority of imported goods. Our estimations strategy employs a three-step iterative estimation process. The first step provides initial estimates of the annual tariffs. The second step refines the estimates by removing observations that exhibit inconsistencies plausibly related to the data entry and rate-switching measurement error discussed above. The third step further improves these estimates by jointly re-estimating the tariffs with data pooled across multiple years. To do so, we jointly estimate the tariffs of a five-digit good over "spells" of consecutively observed years where the variation in the good's tariff is sufficiently small. This final step allows us to smooth estimation error at the annual frequency that could otherwise lead us to mistakenly infer minor fluctuations in legislated tariffs.

In step one, we estimate the following equation separately by product, rate provision code, and year between 1974 and 1988:

$$TD_{vgrt} = \tau_{grt} DV_{vgrt} + f_{grt} Q_{vgrt} + \epsilon_{vgrt}. \tag{5}$$

Naturally, we impose we impose $\tau_{grt} = 0$ for goods only facing specific tariffs and f_{grt}=0 for goods only facing ad valorem tariffs. If two units of quantity are provided, we estimate equation (5) separately for each quantity. If estimated tariffs under each of these quantities provide valid estimates – i.e. $\hat{f}_{grt}, \hat{\tau}_{grt} > 0$ – we use the estimates that give with the lower R^2.[17]

After obtaining initial tariff estimates $\hat{\tau}_{grt}$ and $\hat{f}_{grt}$ for all five-digit rate provision codes, we proceed to the second estimation step. Measurement error for individual observations

[16] This additionally requires that the percentage difference in the quantity amount across the two exporter-port observations is different for each of the two units of quantity. This is satisfied in general as long as the units in the data are not scalars of each other – e.g., pounds and ounces.

[17] In some infrequent cases goods receive numerous specific duties on components of the individual good. Because the IDB measures at most two distinct quantities, neither digitization nor estimation can recover the true tariff on such goods. Due to the infrequency of this type of good and a concern of over-fitting, we do not estimate equation (5) with two units of quantity simultaneously.

is occasionally quite large, such as when countries shift from Column 2 to Column 1 tariffs during the year. We can thus improve the accuracy of our tariff estimates by iteratively removing observations with substantial measurement error.

To this end, we begin by defining ad valorem equivalent tariffs in the data, AVE_{vgrt}, and our estimates of these, $\widehat{AVE}_{vgrt}$:

$$AVE_{vgrt} \equiv \frac{TD_{vgrt}}{DV_{vgrt}} \text{ and } \widehat{AVE}_{vgrt} \equiv \hat{\tau}_{grt} + \frac{\hat{f}_{grt}}{p^*_{vgrt}} \tag{6}$$

where $p^*_{vgrt} = \frac{V_{vgrt}}{Q_{vgrt}}$ is the unit value on the total value of imports V_{vgrt}. For small ad valorem equivalent tariffs, the log estimation error

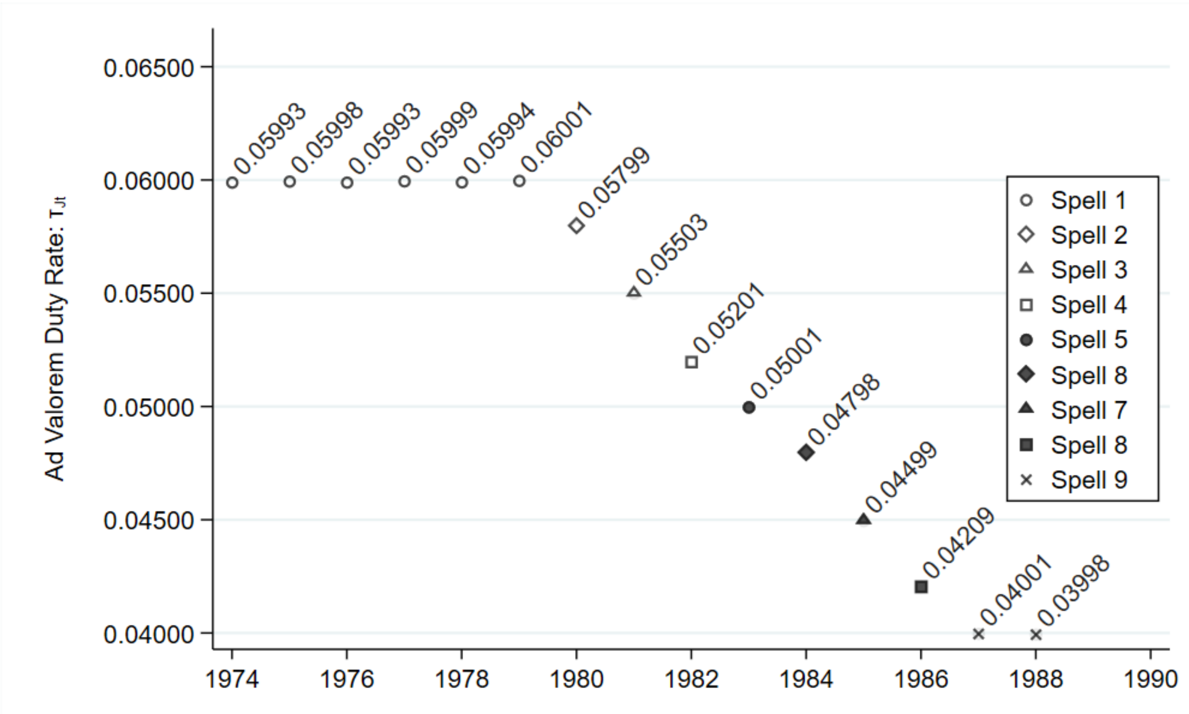
$$\varepsilon^{Est}_{vgrt} \equiv \ln\left[p^*_{vgrt}\left(1 + AVE_{vgrt}\right)\right] - \ln\left[p^*_{vgrt}\left(1 + \widehat{AVE}_{vgrt}\right)\right] \tag{7}$$

approximates the percentage difference between the tariff-inclusive price in the raw data and the value implied by our estimates of the underlying legislated tariffs. In this step we remove variety-level (v) observations with the largest estimation error within each cluster of good-by-rate-provision-by-year (grt) observations. For products where the mean absolute error $|\varepsilon^{Est}_{vgrt}|$ exceeds 0.001, we identify the single observation with the largest standardized error and drop it from our estimating sample.

Using the resulting reduced sample, we repeat the above two steps, again estimating tariffs and removing observations with the largest error within clusters. We repeat this iterative process ten times. If there are still products for which observations exhibit substantial estimation error, we assume that this may be driven by individual countries which have changed MFN status during the year. This would result in systematically high errors in all observations from this country and potentially impact all products exported by the country. Consequently, we drop the set of exporters in a given year for which the mean absolute error $|\varepsilon_{vt}|$ across all variety-by-rate-provision observations within a cluster of export-year observations exceeds 0.001. Finally, we re-estimate the tariffs on the resulting sub-sample to yield our final estimates and iterate on this process up to 10 additional times.

Due to estimation error, our annual estimates often exhibit minor fluctuations around a common mean, suggesting that rates were in fact static during this period. For example, in Figure 2 we can see that TSUSA 11137 had nine distinct tariff rates between 1974 and 1988. The rate was static from 1974 to 1979, was reduced annually from 1980 to 1987, and was constant again between 1987 and 1988. We capture this relationship by identifying spells in the data in which estimated tariffs change by less than 0.015 log points in any two years.

Figure 2: Identifying Tariff Spells from Annual Estimates



Notes: Figure displays annual estimates of ad valorem tariff rates for TSUSA 11137 (Herring, salted or pickled, but not otherwise preserved, and not in airtight containers with weight of no more than 15 lbs) after completing the first two steps of our tariff estimation procedure.

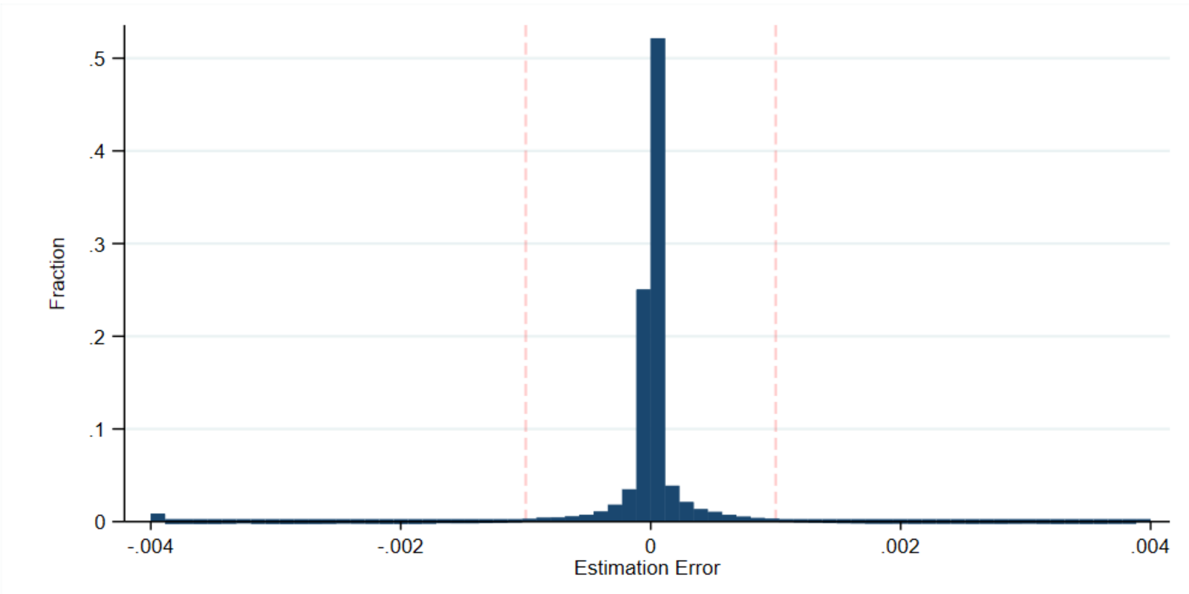
We then repeat steps one and two in our tariff estimation algorithm, pooling observations across time within spells. Again, we eliminate variety-by-rate provision-by-year observations sequentially if they exhibit substantial estimation error and re-estimate tariffs on the reduced sample, and further refine our estimates by eliminating exporter-year pairs which similarly exhibit high average error rates.

2.2 Accuracy and Coverage of Estimated Tariffs

In general, we expect that estimation error will drive a wedge between our tariff estimates and the true legislated values. To assess the extent of this error, we begin by displaying the estimation error ε_{vgrt}, as described in (7), in Figure 3. The dashed red lines indicate thresholds of $|\varepsilon_{vgrt}| = 0.001$, which represents an estimation error of approximately 0.1%

of the tariff-inclusive price. Less than 3.8% of observations lie outside of these thresholds. That is, import prices implied by our estimates very closely match those in the data.

Figure 3: Distribution of Estimation Error



Notes: Estimation error defined in equation (7) at exporter-port-TSUSA 7 digit-rate provision code-year level as log difference between tariff inclusive price and our estimated tariff inclusive price using our tariff estimates from Section 2.1. See main text for further details.

Despite iteratively dropping observations and countries from our sample during tariff estimation, this process yields a database of 7366 unique five-digit TSUSA goods. We restrict our attention to imports for which $|\varepsilon_{vrgt}| < .001$. Our final sample covers nearly 97% of the value of US imports annually. Moreover, we are able to provide disaggregate tariff rates for goods in 1972 and 1973, years in which the IDB does not report duties collected, using our estimated tariff rates from 1974. The notable exception to the high accuracy and coverage of our approach is 1980 where coverage falls to 82%. This is due to two intra-year trade policy changes that render tariff rates incalculable. First, China receives NTR status which means that the country's import and tariff data are subject to both Column I and Column II rates in an indistinguishable manner. Second, tariff changes introduced under the Tokyo round of the GATT are introduced mid-year for some products in this year. Tariff rates for these products tariff rates are similarly incalculable.

Our second method for assessing the accuracy of our estimation strategy is more direct: we digitize all 1978 tariffs using the 1978 TSUSA and compare it with our 1978 estimated tariffs. Collapsing over all varieties, we then calculate log tariff-inclusive price in the data

$\ln(1+AVE^{Data}_{gr,1978})$, using our estimated tariffs, $\ln(1+AVE^{Est.}_{gr,1978})$, and using digitized tariffs $\ln(1+AVE^{Dig.}_{gr,1978})$ This yields a sample of 9422 good-by-rate provision observations. Table 2 reports the correlation between these measures and shows a near-perfect correlation between our estimated tariffs and either those in the data or the digitized tariffs.[18]

Table 2: Sample of Estimated vs Digitized Tariffs

	$\ln(1+AVE^{Data}_{rg,1978})$	$\ln(1+AVE^{Est.}_{rg,1978})$	$\ln(1+AVE^{Dig.}_{rg,1978})$
$\ln(1+AVE^{Data}_{rg,1978})$	1.0000	-	-
$\ln(1+AVE^{Est.}_{rg,1978})$	0.9994	1.0000	-
$\ln(1+AVE^{Dig.}_{rg,1978})$	0.9990	0.9997	1.0000

Notes: Table presents an import weighted pairwise correlation matrix of good-by-rate level tariffs taken from the data, estimated via the algorithm detailed above, and digitized in 1978.

3 US Trade Policy: 1972-1988

With the database of estimated tariffs constructed, we now turn to the behavior of tariff rates during this period, emphasizing two distinct features of the data. First, we investigate the sharp decline in the aggregate tariff rates in the pre-Tokyo Round period despite the absence of any policy change. Second, we provide an explanation for the relative stability of the aggregate AVE in the years following the successful agreement.

3.1 The Accidental Liberalization: 1972-1979

We begin by exploring the unexpected behavior of AVE tariff levels in the first half our sample. As noted above, the aggregate tariff average fell by nearly 50% between 1972 and 1979 despite no apparent changes in legislated tariff rates between 1972 and 1979. Indeed, Section 3.2 will empirically confirm the absence of legislated tariff changes in this period.

To better understand the connection between variety-level price changes and changes in ad valorem equivalent tariff rates, we write the log change in the tariff-inclusive price for

[18]This figure omits the 83 observations (0.8% of our sample in 1978) which exhibit nontrivial discrepancies between data and estimation related to units of quantity in the underlying IDB data.

variety v of good g as

$$\Delta \ln\left(1 + AVE_{vgrt}\right) = \ln\left(1 - \frac{\Delta p^*_{vrgt}}{p^*_{vrgt}} \frac{f_{grt}}{p^*_{vgrt-1}} \frac{1}{1 + AVE_{vgrt-1}}\right) \tag{8}$$

$$= \ln\left(1 - \frac{\Delta p^*_{vgrt}}{p^*_{vgrt}} STS_{vgrt-1} \frac{AVE_{vgrt-1}}{1 + AVE_{vgrt-1}}\right) \tag{9}$$

where

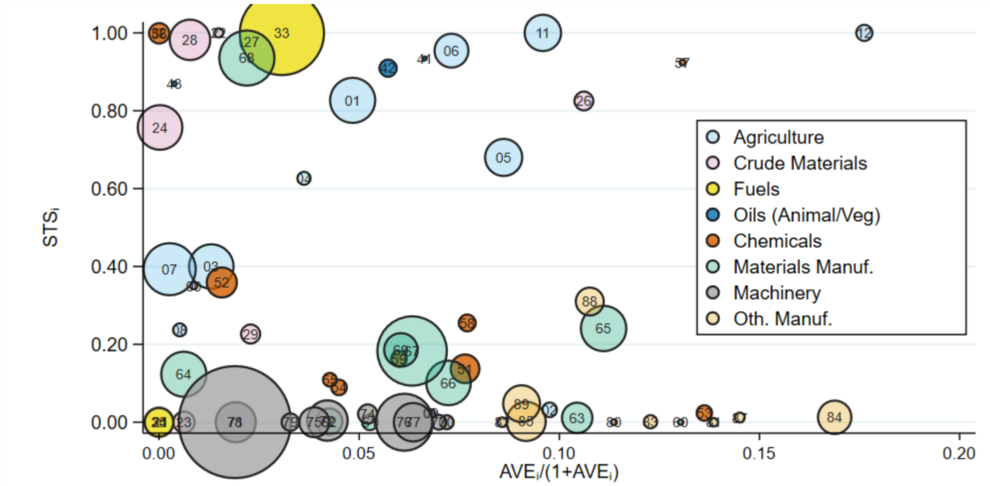
$$STS_{vgrt-1} = \frac{f_{gr}}{p^*_{vgrt}\tau_{grt} + f_{grt}} \tag{10}$$

is the share of total duties accounted for by specific tariffs in year $t-1$ or simply the product's "specific tariff share". As expected, (9) shows that the only source of change in AVE_{vgrt} is changes in variety-level foreign prices $\frac{\Delta p^*_{vgrt}}{p^*_{vgrt}}$ when legislated tariffs remain constant.

Further, (9) shows that exposure to price-driven changes in ad valorem equivalent rates is fully characterized by a product's specific tariff share and its initial AVE level. Intuitively, when tariffs are specified in per-unit terms, their ad valorem equivalent will vary inversely with the price level: rising prices inflate away the protection afforded by specific tariffs, while falling prices enhance it. The products most exposed to price-driven liberalization are thus those with high initial tariff levels that are specified in per-unit terms. Given the substantial inflation during the 1970s and early 1980s, with an average annual change in foreign unit values of approximately 10%, and the pervasiveness of specific tariffs in the era, this is an empirically relevant channel. We quantify the importance of this channel more formally in Section 4, but here outline the extent to which industries were differentially exposed to an "accidental liberalization" through the inflationary erosion of specific tariffs.

To highlight cross-industry variation in this exposure, we construct both specific tariff shares and ad valorem equivalent levels at the two-digit SITC level. Figure 4 documents variation in the two measures across sectors in 1972. Sectors in the upper right of the figure – e.g., Tobacco (12), Plastics in Primary Form (57), Textile Fibers (26), Beverages (11), Sugars & Honey (6), Animal Fats & Oils (41) and Vegetable Fats & Oils (42)) – are those with both high AVE levels and substantial reliance on specific tariffs. It is largely these sectors that are exposed to liberalization in the presence of inflation, and thus drive the decline in aggregate AVE tariffs prior to 1980.

Figure 4: Sectoral Exposure in 1972 to Inflation Erosion of AVE (2-digit SITC)



Notes: Scatter plot of sectoral exposure to intra-policy tariff changes due to inflation/deflation as of 1972. Each bubble is a single 2-digit SITC revision 2 sector whose size reflects import share in 1972. Equation (9) shows exposure to inflation erosion of specific tariff equals product of STS and $\frac{AVE}{1+AVE}$.

3.2 Legislated Tariff Rates

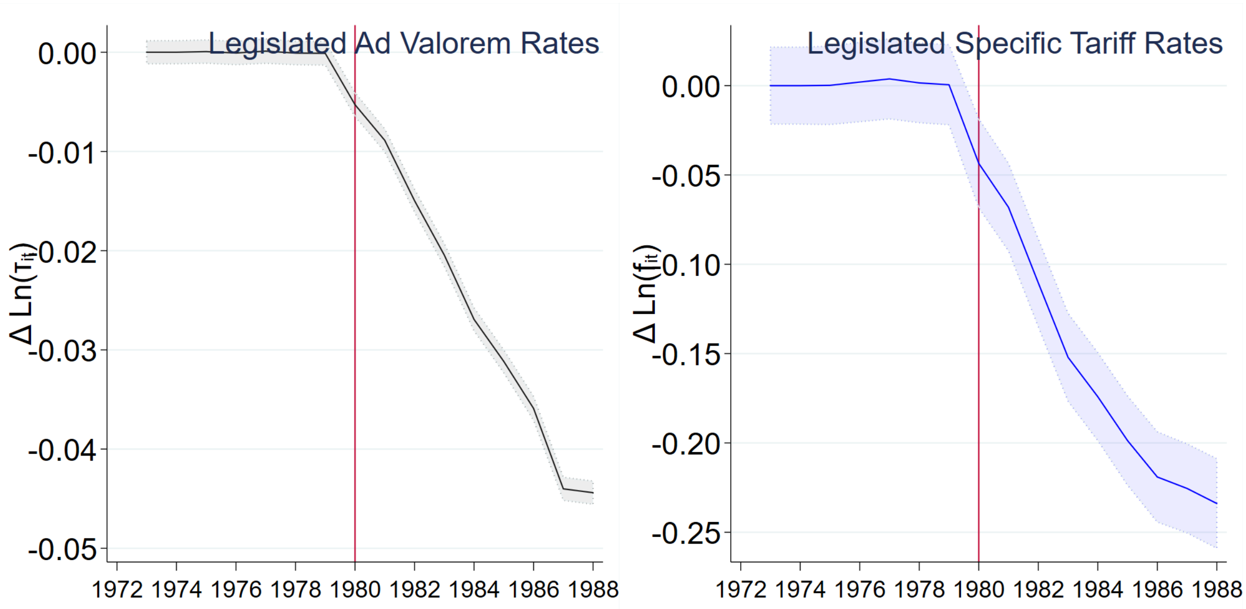
We now turn to the legislated tariff rates. Importantly, our estimated tariffs allow us to distinguish between changes in legislated tariffs and ad valorem equivalent levels, something that is not possible with existing data sets.

Figure 5 examines whether, and if so when, legislated tariffs changed in our sample period 1972-1988. For products with only non-zero ad valorem tariffs, the left panel plots annual point estimates from a regression of $\ln(\tau_{grt})$ on year dummies, with 1972 as the omitted year, and 5-digit TSUSA product fixed effects. That is, the panel plots the average within-product legislated ad valorem tariff relative to the 1972 value. The right panel performs the same exercise for products with only non-zero specific tariffs using $\ln(f_{grt})$.

Two key points emerge from Figure 5. First, the decline in tariff rates during the pre-Tokyo era is not due to a change in legislated US tariffs – legislated rates are indeed static during this era. This confirms the importance of inflation-driven changes in variety-level AVE tariffs as discussed in Section 3.1.

Second, the increase in overall AVE rates during the Tokyo Round phase-outs seen in Figure 1 is not driven by changes in legislated rates. On the contrary, Figure 5 reveals that legislated tariff rates fall substantially once the negotiated tariff cuts begin in 1980. By the

Figure 5: Change in Legislated Tariff Rates 1972-1988



Notes: Figures plot point estimates and 95% confidence intervals from regressions of estimated column 1 legislated ad valorem tariff rates, $\ln(1+\hat{\tau}_{gt})$, in panel (a),$\ln \hat{f}_{gt}$ in panel (b) on time dummies and 5-digit TSUSA fixed effects. See main text for further details.

end of our sample, ad valorem tariff rates fall by 4 percentage points and specific tariff rates fall by by approximately 25% relative to their 1972 baseline.

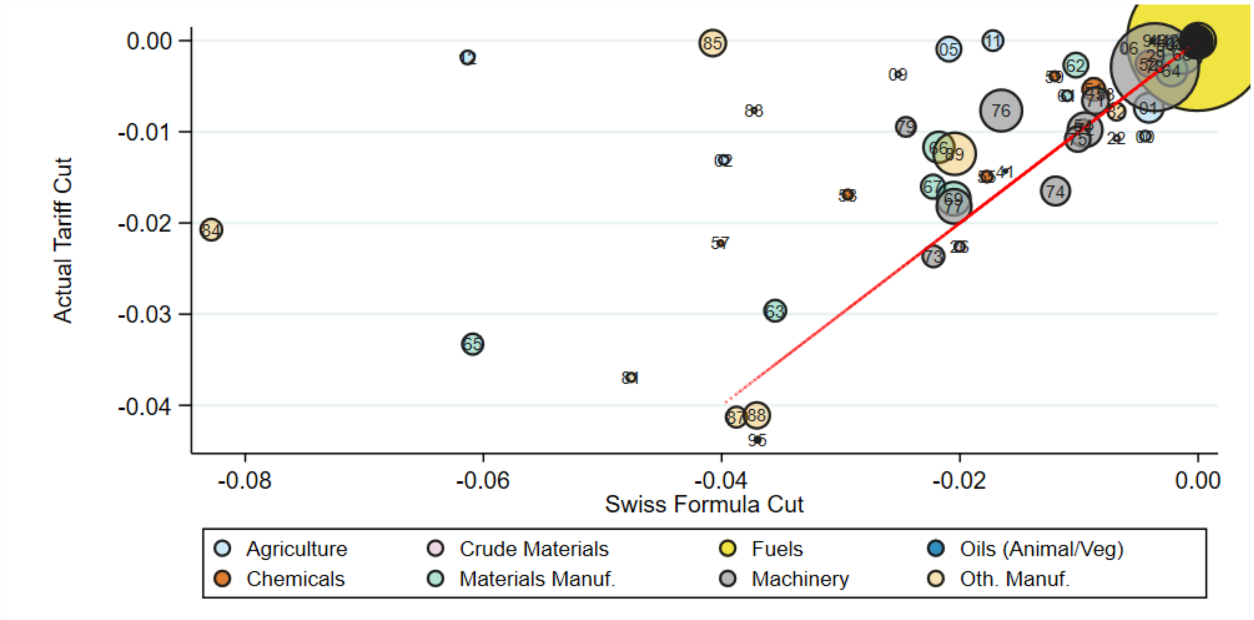
Our data allow us to highlight a separate point concerning the nature of the post-Tokyo tariff cuts. The legislated change in tariffs under the Tokyo Round of the GATT were to be phased in from 1980-1987 according to the "Swiss Formula'. The formula takes an initial rate AVE_{gt_0} and a constant parameter Z that corresponds to the maximum final tariff rate and implies a final rate AVE_{gt} given by

$$AVE_{gt} = \frac{Z \times AVE_{gt_0}}{Z + AVE_{gt_0}} \tag{11}$$

Using import shares and AVE_{gt_0} as of 1979, we can calculate changes in realized tariffs observed under the Tokyo phaseouts and those implied by the Swiss Formula using a value of $Z = 0.14$ (Deardorff and Stern, 1979). Figure 6 illustrates the strong correlation between tariff cuts implied by the Swiss formula and observed tariff cuts. That said, the observed tariff reduction is generally lower and thus consistent with certain products being shielded from the full extent of liberalization. The largest cuts, both in practice and implied by the

Swiss formula, are concentrated in raw materials and materials manufacturing.

Figure 6: Sectoral Exposure in 1979 to Tokyo Phaseouts (2-digit SITC)



These tariff changes are consistent with the standard understanding of GATT negotiations. However, they do not explain the Tokyo-era increase in aggregate AVE rates. To explore this, we turn to a decomposition of changes in aggregate AVE into its constituent pieces. As the aggregate AVE tariff is an import-weighted average of product-level AVE rates, our analysis thus far suggests that changes in the product composition of US imports may play a role in reconciling Figure 1 with Figures 5 and 6.

To explore this possibility, we rewrite the the AVE as

$$
\begin{aligned}
AVE_t &= \frac{\sum_g \sum_v TD_{vgt}}{\sum_g \sum_v M_{vgt}} \\
&= \frac{\sum_g \sum_v M_{vgt}\left(\tau_{vgt} + \frac{f_{vgt}}{p^*_{vgt}}\right)}{\sum_g \sum_i M_{vgt}} \\
&= \sum_g \sum_v s_{vgt}\left(\tau_{vgt} + \frac{f_{vgt}}{p^*_{vgt}}\right) \qquad (12)
\end{aligned}
$$

where AVE_t is the year t sum of total duties across varieties – i.e., exporter by port by seven-digit TSUSA – divided by the sum of variety-level imports M_{vgt}. Similarly, $\tau_{vgt} + \frac{f_{vgt}}{p^*_{vgt}}$

represents the variety-level AVE tariff and s_{vgt} is the variety-level import share. We define three mutually exclusive sets of imported varieties: continuing varieties V^C, exiting varieties V^X, and entering varieties V^N. We can then decompose changes in AVE_t between $t = 0$ and $t = 1$ as follows:

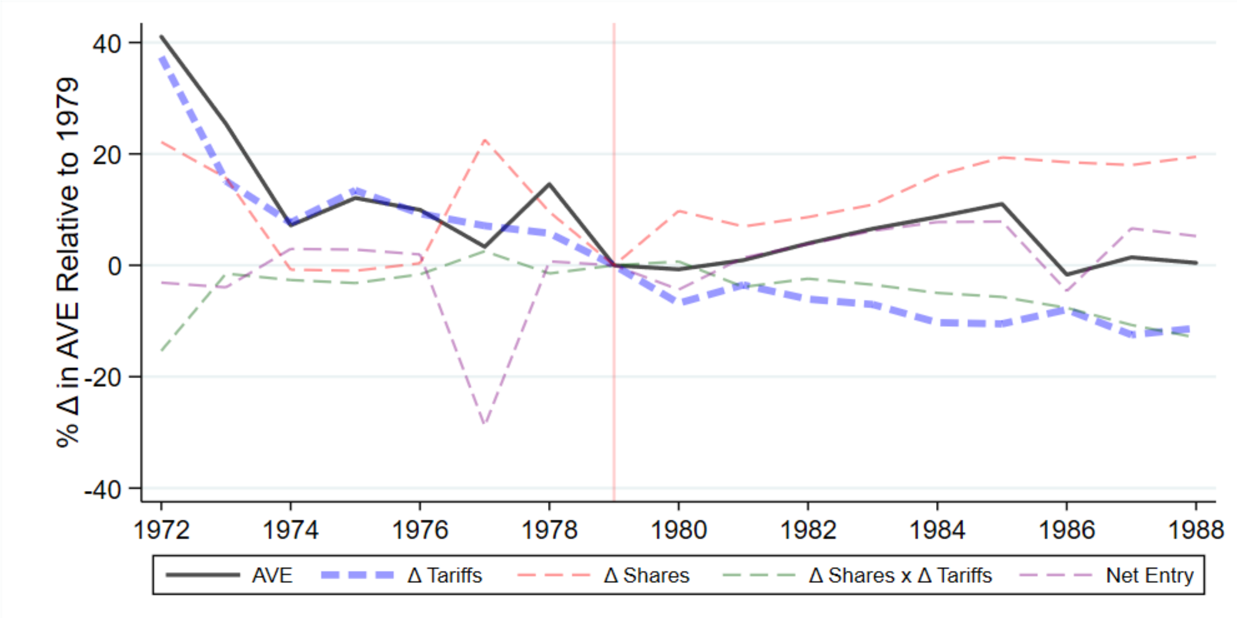
$$\begin{aligned}
\Delta AVE_t \equiv & AVE_{t_1} - AVE_{t_0} \\
= & \underbrace{\sum_{v \in V^C} \Delta s_{vgt} AVE_{vgt_0}}_{\substack{\Delta \text{ Import Shares,} \\ \text{Initial Tariffs Fixed}}} + \underbrace{\sum_{v \in V^C} s_{vt_0} \Delta AVE_{vgt}}_{\substack{\Delta\text{AVE Tariffs,} \\ \text{Initial Shares Fixed}}} + \underbrace{\sum_{v \in V^C} \Delta s_{vgt} \Delta AVE_{vgt}}_{\substack{\Delta \text{ AVE Tariffs,} \\ \Delta \text{ Import Shares}}} \\
& + \underbrace{\sum_{v \in V^N} s_{vgt_1} AVE_{vgt_1} - \sum_{v \in V^X} s_{vgt_0} AVE_{vgt_0}}_{\text{Net Entry}}
\end{aligned} \tag{13}$$

Equation (13) shows changes in the aggregate AVE tariff rate can be divided into the sum of change in tariffs rates on continuing varieties conditional on initial import shares, the product of changes in tariff rates and import shares on continuing products, and changes due to net entry of products. Using our estimated tariff data and defining the base year as 1979, we plot this decomposition in percentage changes in Figure 7. The dark grey line in the figure depicts the aggregate change in AVE_t throughout our sample. Consistent with the USITC aggregate data, we observe a decline in average tariffs prior to 1980 – the AVE is approximately 1.5 percentage points higher in 1972 than in 1979 – and little to no change thereafter (solid black line). And consistent with Figure 5, the dashed blue line reflects a post-1979 fall in legislated tariffs among continuing products holding the composition of imports fixed.

However, Figure 7 highlights two stark points. First, the dashed blue line shows that the average AVE tariff across goods fell continuously in the years preceding the Tokyo Round agreement while holding the import composition across varieties fixed. Thus, (12) implies the only source of time variation underlying falling falling AVE tariffs in the dashed blue is inflation erosion of specific tariffs. Further, this dashed blue line very closely tracks the solid black line which says inflation erosion of specific tariff essentially explains the entire decline in AVE tariffs in the pre-Tokyo period.

Second, changes in the import composition across goods and varieties explains why the aggregate AVE tariff stays fairly flat, and drifts upward if anything, in the post-Tokyo period. The dashed red line shows that, among products continually imported, imports

Figure 7: Decomposition of Percent Change in Aggregate US AVE Tariff Rate



Notes: Plot based on results from equation (13) where $t_0 = 1979$. Indexed AVE plots ΔAVE_t, other lines plot respective components of equation (13). See main text for further details.

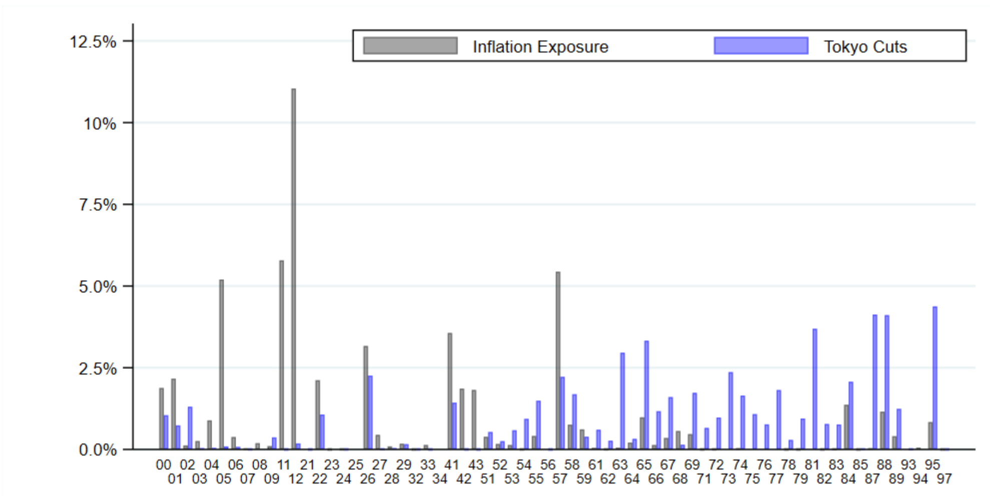
shifted towards products with relatively higher tariff rates. While the dashed green line indicates this composition effect is partially offset by these high-tariff goods receiving larger tariff tariff cuts expanding good higher tariffs goods, the resulting net effect essentially offsets the falling legislated tariffs to leave the aggregate AVE tariff relatively flat.[19]

Ultimately, these results show two distinct liberalizations in the years between 1972 and 1989. The pre-Tokyo years of 1972-1979 saw goods with specific tariffs experience a price-driven liberalization despite the absence of any policy changes. This process continued throughout the 1980s, although lower inflation and the transition of some specific tariffs to ad valorem in the Tokyo Round mitigated the effect of reduced this channel. Following completion of the Tokyo Round in 1979, legislated tariffs declined through the 1980s largely according to the "Swiss Formula", although this was masked in aggregate data by a shift towards higher tariff products.

[19] Perhaps surprisingly, the dashed purple line indicates that the net entry of products (i.e. extensive margin) does provide an offsetting effect to the falling legislated tariffs in terms of explaining the relatively flat aggregate AVE tariff. It is worth noting that we construct products at the five-digit level to match the level at which tariffs are defined. This contrasts with Broda and Weinstein (2006), who document substantial gains from seven-digit TSUSA *varieties*.

To depict industry-level variation in exposure to the two respective liberalizations, Figure 8 illustrates exposure to the price-driven changes in tariff exposure at the two-digit SITC level, as measured by $STS_{it}\frac{AVE_{it}}{1+AVE_{it}}$, as well as the realized tariff cut under the GATT Tokyo. While the negotiated cuts tend to be more pervasive, the accidental tariff reductions are nonetheless substantial with multiple sectors experiencing inflation-driven tariff reductions in excess of 5 percentage points.

Figure 8: Sectoral Exposure to Both Liberalizations (2-digit SITC)



4 Quantifying the Welfare Effects of Specific Tariffs

A primary feature of our data is the ability to separate AVE tariffs into specific and advalorem tariffs. As argued above, the behavior of tariff rates in this era is highly dependent on whether tariffs are specific or advalorem. To formally quantify the aggregate importance of specific tariffs during our sample period, we embed specific tariffs in a standard CES framework and extend existing hat algebra techniques to this setting. In turn, we perform a series of counterfactual scenarios that allow us to isolate and compare the welfare effects of both (i) inflation erosion of specific tariffs and (ii) legislated tariff phase outs mandated by the Tokyo Round of GATT.

4.1 A CES Model with Specific Tariffs

A representative consumer has a three-tier utility function with expenditure E_t allocated at the upper tier across an aggregate domestic composite good q_{Dt} and an aggregate importable composite good q_{Mt} with an elasticity of substitution between these composite goods $\kappa > 1$:

$$u\left(q_{Dt}, q_{Mt}\right) = \left[q_{Dt}^{(\kappa-1)/\kappa} + q_{Mt}^{(\kappa-1)/\kappa}\right]^{\kappa/(\kappa-1)}. \tag{14}$$

The middle tier defines the aggregate importable composite good q_{Mt} as a CES aggregator across composite importable goods q_{gt} with an elasticity of substitution $\gamma > 1$:

$$q_{Mt} = \left[\sum_{g=1}^{G} q_{gt}^{(\gamma-1)/\gamma}\right]^{\gamma/(\gamma-1)}. \tag{15}$$

Finally, the lower tier defines the composite importable good q_{gt} as a CES aggregator across Armington varieties q_{vgt} of good g with a good-specific elasticity of substitution $\sigma_g > 1$ and a time-varying variety-specific taste shock b_{vgt}:

$$q_{gt} = \left[\sum_{v=1}^{V_g} b_{vgt}^{1/\sigma_g} q_{vgt}^{(\sigma_g-1)/\sigma_g}\right]^{\sigma_g/(\sigma_g-1)}. \tag{16}$$

To maintain tractability, we do not model the domestic production structure underlying P_{Dt}. Additionally, we assume aggregate income does not depend on tariff revenue. Each of these channels can be incorporated in a straightforward way.

4.1.1 Import Demand

Maximizing period t utility from good g subject to a fixed level of expenditure E_{gt} yields variety-level demand for good g

$$q_{vgt} = b_{vgt} p_{vgt}^{-\sigma_g} P_{gt}^{\sigma_g - 1} E_{gt} \tag{17}$$

where the good-g price index P_{gt} aggregates over variety-level tariff-inclusive prices p_{vgt}

$$P_{gt} = \left(\sum_{v=1}^{V_g} b_{vgt} p_{vgt}^{1-\sigma_g} \right)^{1/(1-\sigma_g)} . \tag{18}$$

The tariff-inclusive price paid by domestic consumers p_{vgt} is the product of the tariff-exclusive price received by the exporter p_{vgt}^* and the variety-level AVE tariff wedge AVE_{vgt}:

$$p_{vgt} = p_{vgt}^* \left(1 + AVE_{vgt}\right) \tag{19}$$

$$AVE_{vgt} \equiv \tau_{gt} + \frac{f_{gt}}{p_{vgt}^*} \tag{20}$$

with the AVE depending on the variety-level tariff-exclusive price p_{vgt}^*.

Similarly, demand for the good-level composite importable, the aggregate composite importable, and the aggregate composite domestic good are

$$q_{gt} = P_{gt}^{-\gamma} P_{Mt}^{\gamma-1} E_{Mt} \tag{21}$$

$$q_{Mt} = P_{kt}^{-\kappa} P_t^{\kappa-1} E_t \text{ for } k \in \{M, D\} \tag{22}$$

with respective price indices

$$P_{kt} = \left(\sum_{g=1}^{G} P_{gt}^{1-\gamma} \right)^{1/(1-\gamma)} \quad \text{for } k \in \{M, D\} \tag{23}$$

$$P_t = \left(P_{Dt}^{1-\kappa} + P_{Mt}^{1-\kappa} \right)^{1/(1-\kappa)} . \tag{24}$$

We can represent equilibrium demand in terms of expenditure shares. That is,

$$s_{vgt} \equiv \frac{p_{vgt} q_{gt}}{E_{Mt}} = b_{vgt} \left(\frac{P_{gt}}{p_{vgt}} \right)^{\sigma_g - 1} \tag{25}$$

$$s_{gt} \equiv \frac{E_{gt}}{E_{Mt}} = \frac{P_{gt} q_{gt}}{E_{Mt}} = \left(\frac{P_{Mt}}{P_{gt}} \right)^{\gamma-1} \tag{26}$$

$$s_{kt} \equiv \frac{E_{kt}}{E_t} = \frac{P_{kt} q_{kt}}{E_t} = \left(\frac{P_t}{P_{kt}} \right)^{\kappa-1} \text{ for } k \in \{M, D\} . \tag{27}$$

Using (17) and (19), we can measure variety-level expenditure shares in terms of tariff-exclusive prices, $s^*_{vgt} \equiv \frac{p^*_{vgt} q_{vgt}}{E_{gt}}$, and its log form:

$$\text{ln}s^*_{vgt} = -\left(\sigma_g - 1\right)\text{ln}p^*_{vgt} - \sigma_g \text{ln}\left(1 + AVE_{vgt}\right) + \left(\sigma_g - 1\right)\text{ln}P_{gt} + \text{ln}b_{vgt}. \tag{28}$$

The comparative statics of the import demand curve are straightforward. Taste shocks, good-level price indices, and tariffs all shift the import demand curve.

4.1.2 Export supply

Export supply is given by

$$p^*_{vgt} = \exp\left(\eta_{vgt}\right) q^{\omega_g}_{vgt} \tag{29}$$

where η_{gvt} is a time-varying variety-specific technology shock. Taking logs and rewriting in terms of tariff-exclusive variety-level expenditure shares yields

$$\text{ln}s^*_{vgt} = \frac{1 + \omega_g}{\omega_g}\text{ln}p^*_{vgt} - \text{ln}E_{gt} - \frac{1}{\omega_g}\eta_{vgt}. \tag{30}$$

4.1.3 Variety-level equilibrium prices and expenditure shares

Equating import demand (28) and export supply (30) yields the equilibrium tariff-exclusive price

$$\begin{aligned}\text{ln}p^*_{vgt} =& -\frac{\sigma_g \omega_g}{1 + \sigma_g \omega_g}\text{ln}\left(1 + AVE_{vgt}\right) + \frac{\omega_g\left(\sigma_g - 1\right)}{1 + \sigma_g \omega_g}\text{ln}P_{gt} \\ &+ \frac{\omega_g}{1 + \sigma_g \omega_g}\left(\text{ln}E_{gt} + \text{ln}b_{vgt}\right) + \frac{1}{1 + \sigma_g \omega_g}\eta_{vgt}.\end{aligned} \tag{31}$$

and the equilibrium tariff-exclusive expenditure share

$$\begin{aligned}\text{ln}s^*_{vgt} =& -\frac{\sigma_g\left(1 + \omega_g\right)}{1 + \sigma_g \omega_g}\text{ln}\left(1 + AVE_{vgt}\right) + \frac{\left(1 + \omega_g\right)\left(\sigma_g - 1\right)}{1 + \sigma_g \omega_g}\text{ln}P_{gt} \\ &- \frac{\omega_g\left(\sigma_g - 1\right)}{1 + \sigma_g \omega_g}\text{ln}E_{gt} + \frac{1 + \omega_g}{1 + \sigma_g \omega_g}\text{ln}b_{vgt} - \frac{\omega_g\left(\sigma_g - 1\right)}{1 + \sigma_g \omega_g}\eta_{vgt}.\end{aligned} \tag{32}$$

Intuitively, tariffs have two effects on equilibrium prices and expenditure shares and both are mediated through import demand and export supply elasticities: (i) they create a wedge between the tariff-exclusive price received by exporters and the tariff-inclusive prices paid by consumers by decreasing the former and increasing the latter and and (ii) they reduce variety-level expenditure shares for imported goods.

While (31) and (32) characterize the equilibrium tariff-exclusive price and expenditure share, they depend on the endogenous level of good-g expenditure E_{gt}. In turn, E_{gt} depends on relative price indices P_{gt} across importable composite goods g (and hence, e.g., relative tariffs across goods) and the relative price indices P_{kt} for $k = M, D$ across the aggregate importable composite and aggregate domestic composite goods. In turn, the solution to (31) and (32) for a given good g and variety v depend on the variety-level prices of *all* varieties of *all* goods.

Further, the non-linear form of the price indices imply the equilibrium of our model is characterized by a high-dimensional system of non-linear equations. The endogenous variables are (i) the price indices P_t, P_{Mt}, and P_{gt} for $g = 1, ..., G$, (ii) the tariff-exclusive variety-level prices p^*_{vgt} for $v = 1, ..., V_g$ and $g = 1, ..., G$, (iii) the tariff-exclusive variety-level expenditure shares s^*_{vgt} for $v = 1, ..., V_g$ and $g = 1, ..., G$, and (iv) the expenditure levels E_{gt} and E_{Mt}. This is $2V + G + 4$ endogenous variables where $V \equiv \sum_{g=1}^{G} V_g$. Equations (31) and (32) give $2V$ equations. Equation (18) gives G equations. And, equations (23)-(24) and (26)-(27) give 4 equations. The solution to this system for the endogenous variables depends on the following exogenous variables: (i) aggregate expenditure E_t and the aggregate domestic composite good price index P_{Dt}, (ii) legislated tariffs τ_{gt} and f_{vgt}, and (iii) structural import demand and export supply shocks b_{vgt} and η_{vgt}.

Appendix A.1 details how to computationally solve this large non-linear system of equations. This includes how to recover the variety-level unobserved demand shocks which is necessary because our above characterization of the equilibrium treated these demand shocks as known to the modeler. However, a key downside of solving the model in levels is that reliably recovering the variety-level unobserved demand shocks is not straightforward.

To overcome the challenge of recovering the unobserved variety-level demand shocks, the following section characterizes the equilibrium in terms of proportional changes using hat algebra. Formally, this technique performs counterfactuals through a first-order approximation around the equilibrium observed in the data which holds the unobserved demand shocks fixed. While Appendix A.3 shows how to solve our model using exact hat algebra, the exact hat algebra system is high-dimensional, highly non-linear and hence challenging to computationally solve. In contrast, the hat algebra system is linear, only has roughly half the dimensionality of the exact hat algebra system and hence imposes minimial compupational burden. Moreover, the hat algebra solution brings out the key intuition for how specific tariffs work that does not easily emerge in the non-linear exact hat algebra system.

4.2 Hat Algebra

Appendix A.2 contains a complete presentation of our hat algebra system (i.e. $\hat{x} = d\ln x$). Specifically the model's endogenous variables are the proportional changes in (i) the equilibrium price indices $\hat{P}_t$, $\hat{P}_{Mt}$, and $\hat{P}_{gt}$, (ii) the equilibrium tariff-exclusive variety-level prices $\hat{p}^*_{vgt}$, and (iii) the equilibrium expenditure levels $\hat{E}_{Mt}$ and $\hat{E}_{gt}$. These endogenous variables depend on exogenous variables including (i) observed tariffs τ_{vgt} and f_{vgt}, (ii) observed variety-level tariff exclusive prices p^*_{vgt}, (iii) observed expenditure shares s_{Mt}, s_{Dt}, s_{gt} and s_{vgt}, and (iv) estimated import demand and export supply elasticities σ_g and ω_g. The endogenous variables also depend on the following proportional changes in exogenous variables: (i) aggregate expenditure $\hat{E}_t$ (this could be endogenized to include tariff revenue) and the price index for the aggregate domestic composite good $\hat{P}_{Dt}$, (ii) legislated tariffs $\hat{\tau}_{vgt}$ and $\hat{f}_{vgt}$, and (iii) structural import demand and export supply shocks $\hat{b}_{vgt}$ and $\tilde{\eta}_{vgt}$.

Given the complete description of the hat algebra system in Appendix A.2, we use this section to outline the important features that emerge from embedding specific tariffs into an otherwise standard hat algebra system. Indeed, our model and hat algebra approach reduce to a simplified version of Fajgelbaum et al. (2020) in the absence of specific tariffs.

A key observation facilitating our use of hat algebra is that the tariff inclusive price p_{vgt} is multiplicative in the AVE even though it is additive in the specific tariff f_{vgt} (see equations (19)-(20)). Moreover, the proportional change in the AVE is

$$
\begin{aligned}
(\widehat{1 + AVE_{vgt}}) &= \frac{d\tau_{vgt} + \frac{1}{p^*_{vgt}} df_{vgt} - \frac{f_{vgt}}{\left(p^*_{vgt}\right)^2} dp^*_{vgt}}{1 + AVE_{vgt}} \\
&\equiv \widetilde{AVE}_{vgt} - \hat{p}^*_{vgt} STS_{vgt} \frac{AVE_{vgt}}{1 + AVE_{vgt}}
\end{aligned} \tag{33}
$$

where

$$
\widetilde{AVE}_{vgt} = \hat{\tau}_{vgt} \left(1 - STS_{vgt}\right) \frac{AVE_{vgt}}{1 + AVE_{vgt}} + \hat{f}_{vgt} STS_{vgt} \frac{AVE_{vgt}}{1 + AVE_{vgt}}. \tag{34}
$$

Thus, the proportional change in the AVE is additive in proportional changes of two terms: (i) *legislated* tariff changes $\widetilde{AVE}_{vgt}$ and (ii) *inflation erosion* of specific tariffs.

This additive property of $(1+\widehat{AVE}_{vgt})$ intuitively influences two key parts of the hat algebra system. First, specific tariffs affect how changes in tariff-exclusive prices pass through to tariff-inclusive prices:

$$\begin{aligned}\hat{p}_{vgt} &= \hat{p}^*_{vgt} + (1+\widehat{AVE}_{vgt}) \\ &= \left(1 - STS_{vgt}\frac{AVE_{vgt}}{1+AVE_{vgt}}\right)\hat{p}^*_{vgt} + \widetilde{AVE}_{vgt}\end{aligned} \tag{35}$$

Intuitively, increases in tariff-exclusive prices generate smaller pass-through to tariff-inclusive prices in the presence of specific tariffs because the associated inflation erodes the AVE of the specific tariff component in the tariff-inclusive price. That is, a doubling of the tariff-exclusive price leads to a less than doubling of the tariff-inclusive price in the presence of specific tariffs. In the traditional case of no specific tariffs, $STS_{vgt} = 0$ and a doubling of the tariff-exclusive price would also double the tariff-inclusive price. Naturally, the dampened pass-through flows through to good-level (and more aggregate) price indices:

$$\hat{P}_{gt} = \sum_{v=1}^{V_g} s_{vgt}\left(1 - STS_{vgt}\frac{AVE_{vgt}}{1+AVE_{vgt}}\right)\hat{p}^*_{vgt} + \sum_{v=1}^{V_g} s_{vgt}\widetilde{AVE}_{vgt}. \tag{36}$$

Second, specific tariffs also affect the magnitude of tariff-exclusive price changes themselves in response to underlying shocks. Using (33) and imposing that the structural import demand and export supply shocks do not change from the observed data, we have

$$\hat{p}^*_{vgt} = -\frac{\sigma_g\omega_g\varphi^{-1}_{vgt}}{1+\sigma_g\omega_g}\widetilde{AVE}_{vgt} + \frac{\omega_g\varphi^{-1}_{vgt}}{1+\sigma_g\omega_g}\left[\hat{E}_t + (\kappa-1)\hat{P}_t + (\gamma-\kappa)\hat{P}_{Mt} + (\sigma_g-\gamma)\hat{P}_{gt}\right] \tag{37}$$

where

$$\varphi^{-1}_{vgt} \equiv \left(1 + \frac{1}{1+\sigma_g\omega_g}STS_{vgt}\frac{AVE_{vgt}}{1+AVE_{vgt}}\right)^{-1} \leq 1. \tag{38}$$

In the special case of no specific tariffs, the expression for $\hat{p}^*_{vgt}$ in (37) reduces to that found in the existing literature because $STS_{vgt} = 0$ implies $\varphi_{vgt} = 1$. More generally, $\varphi^{-1}_{vgt} < 1$ in the presence of specific tariff captures the intuition that the increase in p^*_{vgt} due to lower legislated tariffs (or, e.g., higher aggregate expenditure) is smaller in the presence of specific tariffs because the higher p^*_{vgt} erodes the AVE of the specific tariff.

Importantly, these two effects reinforce each other. Not only do specific tariffs dampen the pass-through of any given tariff-exclusive price change to the tariff-inclusive price. They also mitigate the magnitude of tariff-exclusive price changes in response to underlying economic shocks.

4.3 Import Demand and Export Supply Elasticities

To separate the role of AVE tariff changes driven by legislated tariff changes and price movements, the hat algebra system requires import demand and export supply elasticities. Given the absence of obvious instruments to serve as demand and supply shifters, we generalize the methodology of Feenstra (1994) and Broda and Weinstein (2006) (henceforth F/BW) to a setting with specific tariffs.[20]

After differencing the import demand and export supply equations (28) and (30) both over time within a variety and relative to a reference variety k (i.e. $\Delta^k x_{vgt} \equiv (x_{vgt} - x_{vg,t-1}) - (x_{kgt} - x_{kg,t-1})$), we multiply the resulting double-differenced import demand and export supply equations to obtain a generalization of the Feenstra (1994) estimating equation:

$$\begin{aligned}\left(\Delta^k \ln p^*_{vgt}\right)^2 =& \theta_1 \left(\Delta^k \ln s^*_{vgt}\right)^2 + \theta_2 \Delta^k \ln p^*_{vgt} \Delta^k \ln s^*_{vgt} \\ &+ \theta_3 \Delta^k \ln p^*_{vgt} \Delta^k \ln\left(1 + AVE_{vgt}\right) + \theta_4 \Delta^k \ln s^*_{vgt} \Delta^k \ln\left(1 + AVE_{vgt}\right) + u_{vgt}\end{aligned} \tag{39}$$

where $u_{vgt} \equiv \frac{\varepsilon^k_{vgt}\delta^k_{vgt}}{\sigma-1}$, $\theta_1 = \frac{w}{(1+w)(\sigma-1)}$, $\theta_2 = -\frac{w(2-\sigma)+1}{(1+w)(\sigma-1)}$, $\theta_3 = \frac{\sigma}{\sigma-1}$, and $\theta_4 = \frac{\sigma w}{(1+w)(\sigma-1)}$. In the special case of $\Delta^k \ln\left(1 + AVE_{vgt}\right) = 0$, this equation exactly reduces to that used in F/BW.

Implicitly assuming $\Delta^k \ln\left(1 + AVE_{vgt}\right) = 0$, Feenstra (1994) shows that θ_1 and θ_2 can be consistently estimated using variety dummies as instruments in an IV approach or, equivalently, by time-demeaning the data and using weighted least squares with one observation per variety and weights corresponding to the number of variety-level observations. In our context, this corresponds to observations at the exporter-by-5 digit TSUSA level, with the number of years as weights.[21] In turn, one can use the theoretically implied relationship

[20] In their study of the US-China trade war, Fajgelbaum et al. (2020) estimate these elasticities using the method proposed in Zoutman et al. (2018). To do so, they rely on the discriminatory nature of US trade war tariffs on China. We cannot follow such an approach because changes in tariffs that we study impact nearly all US trading partners identically.

[21] With two endogenous explanatory variables, one needs at least two instruments for identification of the relevant parameters. Using variety dummies as instruments, this requires at least two varieties in addition to the reference variety. With four endogenous explanatory variables in (39), one needs four varieties in addition to the reference variety. Adding a constant to either specification comes with the requirement of an additional instrument, and hence an additional variety.

between the parameter estimates $\hat{\theta}_1$ and $\hat{\theta}_2$ and σ and ω to recover the structural elasticities.

However, if $\Delta^k \ln(1 + AVE_{vgt}) \neq 0$ then the F/BW estimating equation produces biased estimates of θ_1 and θ_2, and thus of σ_g and ω. This is the empirically relevant case in our setting because

$$\begin{aligned}\Delta^k \ln(1 + AVE_{vgt}) = &\left[\ln\left(1 + \tau_{vgt} + \frac{f_{vgt}}{p^*_{vgt}}\right) - \ln\left(1 + \tau_{vgt-1} + \frac{f_{jt-1}}{p^*_{vgt-1}}\right)\right] \\ &- \left[\ln\left(1 + \tau_{kgt} + \frac{f_{kgt}}{p^*_{kgt}}\right) - \ln\left(1 + \tau_{kgt-1} + \frac{f_{kgt-1}}{p^*_{kgt-1}}\right)\right]\end{aligned}$$

is generally non-zero in the presence of specific tariffs. In particular, differential changes in foreign prices across varieties will generate differential changes in ad valorem equivalent tariffs *even if* legislated tariffs τ_{gt} and f_{gt} are constant both across exporters and over time.[22] Further, the interaction terms corresponding to θ_3 and θ_4 in (39), which are omitted from the standard F/BW approach, will generally be correlated with those for θ_1 and θ_2. Hence, omitting these terms will produce biased estimates in the presence of specific tariffs.

Naturally, the magnitude of bias is an empirical question. To address this, we estimate (39) both with and without the θ_3 and θ_4 terms, using weighted non-linear least squares on time-demeaned data. Constrained non-linear least squares allows us to directly estimate the structural elasticities σ_g and ω while respecting the constraints imposed on σ_g and ω by the reduced form parameters θ_1, θ_2, θ_3 and θ_4 in (39). We discuss our estimates and quantify the biases associated with the canonical Feenstra (1994) estimator in Appendix B.

4.4 Counterfactual Results

As detailed in Section 4.2 and Appendix A.2, we can represent the proportional change in equilibrium outcomes for all endogenous variables as a function of changes in tariffs while holding constant unobserved import demand and export supply shocks as well as aggregate expenditure and the domestic price index. To quantify the importance that specific tariffs played in the welfare gains from trade liberliazation in this era, we consider two separate counterfactuals which we compare to the benchmark equilibrium observed in the data. By

[22] In the absence of specific tariffs, $\Delta^k \ln(1 + AVE_{vgt}) = 0$ will hold when all exporters of the good face the same tariff (e.g. an MFN tariff) even if this tariff changes over time. However, $\Delta^k \ln(1 + AVE_{vgt}) = 0$ will not hold when some exporters face the MFN tariff but other countries face non-MFN tariffs (due to, e.g. Free Trade Agreements or GSP-type programs) and this preferential tariff margin changes over time.

definition, these benchmark data incorporate both the inflationary erosion of specific tariffs and the legislated GATT Tokyo Round tariff phase-outs.

The first counterfactual considers an environment where all tariffs are advalorem tariffs and there is no tariff liberalization throughout the entire period 1972-1988. This will reveal the overall welfare gains from the legislated tariff cuts mandated by the Tokyo Round and inflationary erosion of specific tariffs. To do so, we fix all tariffs at an ad-valorem equivalent rate defined by the good's tariff(s) in 1972 (or the earliest year it was observed) and leave them unchanged thereafter. For goods protected by specific or compound tariffs, we set the price of each variety within good g equal to the mean price p^*_{vgt} across all varieties of good g in 1972 when converting into an ad valorem equivalent tariff. Given these fixed rates, our counterfactual analysis calculates the counterfactual aggregate importable price index P_{Mt}, i.e. without any specific tariffs and without any tariff liberalization, *relative* to that in the observed data.[23]

The second counterfactual considers an environment where all tariffs are ad valorem but legislated tariffs get phased out as mandated by the Tokyo Round GATT negotiations. In other words, this counterfactual only removes inflationary erosion of specific tariffs from the data; it does not remove legislated tariff cuts from the data. As such, this counterfactual will reveal the welfare gains from inflationary erosion of specific tariffs and the difference between the two counterfactuals reveals the welfare gains from the Tokyo Round legislated tariff cuts. For goods with specific or compound duties, we set the time-varying ad valorem tariff as the ad valorem equivalent tariff implied by allowing legislated rates to change as in the data but fixing their unit values p^*_{vgt} at their 1972 level as in the first counterfactual. Again, this second counterfactual analysis calculates the counterfactual aggregate importable price index P_{Mt}, i.e. with legislated tariff cuts but without inflationary erosion of specific tariffs, relative to that in the observed data.

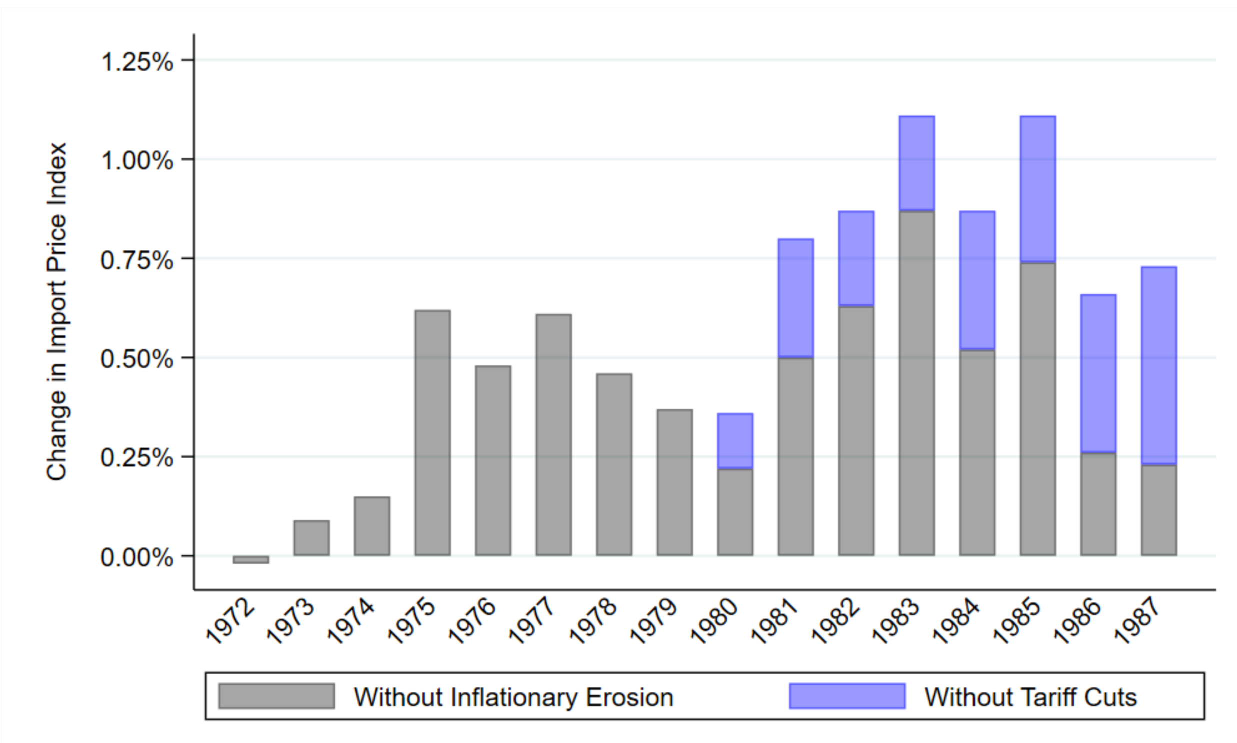
Figure 9 presents the counterfactual results. The height of each bar is the total change in the aggregate importable price index P_{Mt} if all liberalization were shut down – that is, the change implied by the first counterfactual. The gray portion of each bar corresponds to the share attributable to the inflationary erosion of specific tariffs – that is, the change implied by the second counterfactual. The residual is thus the component driven by changes

[23]Note that compensating variation as a share of initial aggregate expenditure due to a tariff-induced change in the aggregate price index P_t is $\Delta P_t/P_t \approx \hat{P}_t$ where $\hat{P}_t = s_{Mt}\hat{P}_{Mt} + s_{Dt}\hat{P}_{Dt}$. That is, together with observed data and an assumption about the change in the domestic price index, the proportional change in the price index for the aggregate import composite good $\hat{P}_{Mt}$ is sufficient to characterize the change in welfare.

in legislated tariffs.

Figure 9 starkly illustrates that the "accidental liberalization" plays a dominant role. Every year after 1972, the conversion to a purely ad valorem tariff schedule would have caused an increase in the import price index, with a magnitude that tends to increase throughout the sample as domestic inflation leads to more erosion of tariffs over time. This effect averages a 0.42% increase annually for 17 years and it peaks at at a 0.87% increase in 1983. By contrast, the removal of the Tokyo Round phaseouts only impacts tariff rates from 1980-1988 and yields an average 0.32% increase in importer prices for each of those 9 years. While both channels have a meaningful impact on consumer welfare, inflationary erosion of specific tariffs is the more important determinant of import prices and consumer welfare during our sample. As our results above, this suggests that the choice to focus solely on AVE levels and ignoring the distinction between ad valorem and specific tariffs in this era is not innocuous.

Figure 9: Counterfactual Change in Importer Price Index



5 Conclusion

Despite having a tariff liberalization larger in magnitude than all combined liberalization that followed, the decades immediately before and after the Tokyo Round of the GATT have received relatively little empirical attention. In this paper, we construct the first dataset on annual legislated tariffs between 1972 and 1988 that we hope will prove a valuable resource for subsequent research.

With this new tariff-line level data, we explore the evolution of US tariffs in this era. Importantly, specific tariffs account for between 25%-35% of US tariff lines during this era. While typically overlooked, we show that specific tariffs play a critical role in shaping the evolution of aggregate ad valorem tariff rates and import price indices. Rampant inflation in the era could naturally erode the protective capacity of specific tariffs and substantially lower both import prices and consumer welfare.

To quantify these impacts we embed specific and ad valorem tariffs into an otherwise standard CES structure and extend existing hat algebra techniques to this setting so we can perform counterfactual analyses. In doing so, we also extend the work of Feenstra (1994); Broda and Weinstein (2006) to estimate model-consistent import demand and supply elasticities. We show that the US underwent an *accidental* liberalization over the period 1972-1979 immediately prior to implementing agreed upon tariff reductions under the Tokyo Round of the GATT. Our counterfactuals show the inflationary erosion of specific tariffs had a greater impact on impoirt price indices and consumer welfare than the legislated Tokyo Round tariff cuts.

Our analysis thus far leaves many avenues for subsequent research. The dramatic "accidental" liberalization of US ad valorem equivalent tariffs through inflationary foreign price shocks should have notable impacts on US import growth, US output, and US labor markets. Finally, the per unit nature of specific tariffs and existing literature linking quality to per unit trade costs leads to interesting issues involving endogenous quality.

References

Ahmad, J. (1978). Tokyo rounds of trade negotiations and the generalised system of preferences. *The Economic Journal 88*(350), 285–295. 4

Anderson, J. E. and E. Van Wincoop (2004). Trade costs. *Journal of Economic literature 42*(3), 691–751. 1

Autor, D., D. Dorn, and G. H. Hanson (2013). The China Syndrome: Local Labor Market Effects of Import Competition in the United States. *The American Economic Review 103*(6), 2121–2168. 1

Boehm, C. E., A. A. Levchenko, and N. Pandalai-Nayar (2020). The long and short (run) of trade elasticities. *NBER w27064*. 1

Bond, E. W., M. J. Crucini, T. Potter, and J. Rodrigue (2013). Misallocation and productivity effects of the smoot-hawley tariff. *Review of Economic Dynamics*, 120–134. 1

Broda, C. and D. E. Weinstein (2006). Globalization and the gains from variety. *The Quarterly Journal of Economics*. 1, 19, 4.3, 5

Brown, F. and J. Whalley (1980). General equilibrium evaluations of tariff-cutting proposals in the tokyo round and comparisons with more extensive liberalisation of world trade. *The Economic Journal 90*(360), 838–866. 4

Caliendo, L. and F. Parro (2015). Estimates of the trade and welfare effects of nafta. *Review of Economic Studies 82*(1'). 1

Caliendo, L. and F. Parro (2021). Trade policy. *NBER w. 29051*. 1

Chan, K. S. (1985). The international negotiation game: some evidence from the tokyo round. *The Review of Economics and Statistics*, 456–464. 4

Cox, L. (2022). The long-term impact of steel tariffs on u.s. manufacturing. *working paper*. 1

Crucini, M. J. (1994). Sources of variation in real tariff rates: The united states, 1900-1940. *American Economic Review 84*(3). 1

Deardorff, A. V. and R. M. Stern (1979). *An Economic Analysis of the Effects of the Tokyo Round of Multilateral Trade Negotiations on the United States and the Other Major Industrialized Countries: A Report*, Volume 5. US Government Printing Office. 3.2

Deardorff, A. V. and R. M. Stern (1981). A disaggregated model of world production and trade: An estimate of the impact of the tokyo round. *Journal of Policy Modeling 3*(2), 127–152. 4

Deardorff, A. V. and R. M. Stern (1983). Economic effects of the tokyo round. *Southern Economic Journal*, 605–624. 4

Fajgelbaum, P. D., P. K. Goldberg, P. J. Kennedy, and A. K. Khandelwal (2020). The return to protectionism. *The Quarterly Journal of Economics 134*(1), 1–55. 1, 4.2, 20, A.2

Feenstra, R. C. (1994). New product varieties and the measurement of international prices. *The American Economic Review 84*(1), 157–177. 1, 4.3, 4.3, 5

Feenstra, R. C. (1996). US imports, 1972-1994: Data and concordances. *NBER WP No. 5515*. 1, 2, 11

Feenstra, R. C. and J. Romalis (2014). International prices and endogenous quality. *The Quarterly Journal of Economics*. 1

Flaaen, A. and J. Pierce (2021). Disentangling the effects of the 2018-2019 tariffs on a globally connected u.s. manufacturing sector. *Federal Reserve Board Finance and Economics Discussion*. 1

Goldberg, P. and N. Pavcnik (2016). The effects of trade policy. *Handbook of Commercial Policy*. 1

Greenland, A. and J. Lopresti (2022). Trade policy as an exogenous shock: Focusing on the specifics. *Upjohn Institute working paper*. 1

Greenland, A., J. Lopresti, and P. McHenry (2019). Import competition and internal migration. *Review of Economics and Statistics 101*(1), 44–59. 1

Hakobyan, S. and J. McLaren (2016). Looking for Local Labor Market Effects of NAFTA. *The Review of Economics and Statistics 98*(4), 728–741. 1

Handley, K. and N. Limao (2017, September). Policy uncertainty, trade, and welfare: Theory and evidence for china and the united states. *American Economic Review 107*(9), 2731–83. 1

Hummels, D. and A. Skiba (2004). Shipping the good apples out? an empirical confirmation of the alchian-allen conjecture. *Journal of Political Economy 112*(6), 1384–1402. 1

Irarrazabal, A., A. Moxnes, and L. D. Opromolla (2015). The tip of the iceberg: A quantitative framework for estimating trade costs. *Review of Economics and Statistics 97*, 777–792. 1

Irwin, D. A. (1998). Changes in us tariffs: The role of import prices and commercial policies. *American Economic Review 88*(4). 1

Kovak, B. K. and P. M. Morrow (2022). The long-run labor market effects of the canada-u.s. free trade agreement. *NBER w. 29793*. 1

Lake, J. and D. Liu (2021). Local labor market effects of the 2002 bush steel tariffs. *working paper*. 1

Pierce, J. R. and P. K. Schott (2016). The Surprisingly Swift Decline of U.S. Manufacturing Employment. *American Economic Review 106*(7), 1632–1662. 1

Sato, K. (1976). The ideal log-change index number. *The Review of Economics and Statistics*, 223–228. A.3

Soderbery, A. (2015). Estimating import supply and demand elasticities: Analysis and implications. *The Journal of International Economics 96*, 157–177. 1

Sørensen, A. (2014). Additive versus multiplicative trade costs and the gains from trade liberalizations. *Canadian Journal of Economics/Revue canadienne d'Ã©conomique 47*, 1032–1046. 1

Trefler, D. (1993). Trade liberalization and the theory of endogenous protection: An econometric study of u.s. import policy. *Journal of Political Economy 101*. 1

Van Cott, T. N. and L. J. Wipf (1983). Tariff reduction via inflation: Us specific tariffs 1972–1979. *Review of World Economics 119*(4), 724–733. 5

Vartia, Y. O. (1976). Ideal log-change index numbers. *scandinavian Journal of statistics*, 121–126. A.3

Zoutman, F., Gavrilova, and A. O. Hopland (2018). Estimating both supply and demand elasticities using variation in a single tax rate. *Econometrica*. 20

A Counterfactual Analysis

A.1 Solving the Model in Levels

Two computational issues prove challenging for solving our model in levels. The first issue is that the system of equations characterizing the equilibrium of our model – described in Section 4.1.3 – is both high-dimensional and highly non-linear. The second issue is recovering the unobserved variety-level structural demand shocks b_{vgt} and structural supply shocks η_{vgt} so that one can not only hold them constant in the counterfactuals but also so one can compute the price indices P_{gt} which depend on the demand shocks b_{vgt}). While recovering the supply shocks is straightforward, reliably recovering the demand shocks is challenging.

These issues motivate our use of hat algebra to derive proportional changes in equilibrium variables under our counterfactual analysis rather than solving for the level of the counterfactual equilibrium variables. Nevertheless, Appendices A.1.1 and A.1.2 show how to solve these two issues.

A.1.1 An Algorithm for Solving the Model in Levels

Throughout this subsection, suppose that we have recovered the unobserved structural import demand and export supply shocks (Appendix A.1.2 details how they can be recovered). The algorithm then has an inner, middle and upper nest corresponding to the different nests of the utility function.

Inner Nest

1. Take values of b_{vgt} for $v = 1, ..., V_g$ and $g = 1, ..., G$ and σ_g for $g = 1, ..., G$ as given. Pick initial values for good-level expenditures E_{gt} for $g = 1, ..., G$ and tariff-exclusive prices p^*_{vgt} for $v = 1, ..., V_g$ and $g = 1, ..., G$.

2. Calculate good-level price indices P_{gt} for $g = 1, ..., G$ using (18).

3. Calculate variety-level demand-side and supply-side expenditure shares $\ln s^*_{vgt}$ for $v = 1, ..., V_g$ and $g = 1, ..., G$ using (28) and (30).

4. If excess supply (demand) for variety v of good g in year t then scale p^*_{vgt} down (up) by a factor of $1 - x$ $(1 + x)$ where $x \in (0, 1)$.

5. Repeat steps 2-4 of the inner nest until convergence.

Middle Nest

1. Take value of γ as given. Pick initial value of importable expenditure E_{Mt}. Combining E_{Mt} with the initial values of good-level expenditures E_{gt} in step 1 of the inner nest gives initial log values of good-level tariff-inclusive expenditure shares $\ln s_{gt} = \ln\frac{E_{gt}}{E_{Mt}}$ for $g = 1, ..., G$.

2. Calculate aggregate importable price index P_{Mt} using (23) and the good-level price indices P_{gt} from the inner nest.

3. Calculate good-level expenditure shares $\ln s_{gt}$ using of (26): $\ln s_{gt} = \ln\left(\frac{P_{gt}}{P_{Mt}}^{(\gamma-1)}\right)$

4. If calculated good-level expenditure share $\ln s_{gt}$ from step 3 is higher (lower) than its initial value from step 1 then scale calculated expenditure share $\ln s_{gt}$ down (up) by a factor of $1 - e$ $(1 + e)$ where $e \in (0, 1)$.

5. Given scaled values of $\ln s_{vgt}$ for $g = 1, ..., G$ and given value of E_{Mt}, use implied values of E_{gt} for $g = 1, ..., G$ as initial values in step 1 of the inner nest.

6. Repeat steps 1-4 of the middle nest.

7. Repeat steps 5-6 of the middle nest until convergence.

Upper Nest

1. Take values of E_t, P_{Dt}, and κ as given. Combining E_t with the initial value of E_{Mt} in step 1 of the middle nest gives an initial value of the aggregate importable expenditure share $s_{Mt} = \frac{E_{Mt}}{E_t}$.

2. Calculate aggregate price index P_t using (24) and P_{Mt} from the middle nest.

3. Calculate expenditure share s_{Mt} using (27): $s_{Mt} = \frac{P_{Mt}}{P_t}^{(\kappa-1)}$

4. If calculated aggregate importable expenditure share s_{Mt} from step 3 is higher (lower) than its initial value from step 1 then scale calculated expenditure share down (up) by a factor of $1 - z$ $(1 + z)$ where $z \in (0, 1)$.

5. Given the scaled value of s_{Mt} and the value of E_{Mt}, use implied value of E_{Mt} as initial values in step 1 of the middle nest.

6. Repeat steps 1-4 of the upper nest.

7. Repeat steps 5-6 of the upper nest until convergence.

A.1.2 Recovering Unobserved Structural Demand and Supply Shocks

The key challenge to counterfactual analysis using our model in levels is recovering the unobserved import demand and export supply shocks. On one hand, recovering the unobserved structural supply shocks is straightforward. Manipulating the export supply equation (30) reveals

$$\eta_{vgt} = -\omega_g \ln s^*_{vgt} + (1 + \omega_g) \ln p^*_{vgt} - \omega_g \ln E_{gt} \tag{40}$$

where all of the variables on the right hand side are either observed ($\ln s^*_{vgt}$, $\ln p^*_{vgt}$, and $\ln E_{vgt}$) or estimated (ω_{vgt}).

On the other hand, recovering the unobserved variety-level demand shock for a given variety v of good g in time t is significantly complicated by the fact that these demand shocks affect import demand not only directly but also through the good-level price index and hence on all variety-level prices and all variety-level demand shocks of good g in time t. That is, the following highly non-linear system of V_g simultaneous equations defines the unobserved demand shocks b_{vgt} for good g in time t as a function of variables that are either observed ($\ln s_{vgt}$ and $\ln p_{vgt}$) or estimated (σ_g):

$$\ln s_{vgt} + (\sigma_g - 1) \ln p_{vgt} = (\sigma_g - 1) \ln P_{gt} + \ln b_{vgt} \tag{41}$$

$$\ln P_{gt} = \frac{1}{1-\sigma_g} \ln \left(\sum_{v=1}^{V_g} b_{vgt} p_{vgt}^{1-\sigma_g} \right). \tag{42}$$

Solving this system of non-linear equations is not straightforward.

While still not straightforward, a simpler approach is to focus on recovering relative unobserved demand shocks. Starting at the lower level of utility and using (25), the unobserved demand shock for good g in time t of variety v relative to a *reference variety* $k \neq v$ is

$$\frac{b_{vgt}}{b_{kgt}} = \frac{s_{vgt}}{s_{kgt}} \left(\frac{p_{kgt}}{p_{vgt}}\right)^{\sigma_g - 1} \equiv \alpha_{vgt} \tag{43}$$

where α_{vgt} consists of either variables that are either observed (s_{vgt}, s_{kgt}, p_{vgt} and p_{kgt}) or estimated (σ_g). Because CES prices are homogeneous of degree one, the good-level price index is proportional to the unobserved demand shock of the good's reference variety and a psuedo-price index $\tilde{P}_{gt}$ that only depends on variables either observed (α_{vgt} and p_{vgt}) or estimated (σ_{vgt}):

$$P_{gt} = b_{kgt} \left(\sum_{v=1}^{V_g} \alpha_{vgt} p_{vgt}^{1-\sigma_g}\right)^{\frac{1}{1-\sigma_g}} \equiv b_{kgt} \tilde{P}_{gt} \tag{44}$$

Moving to the middle level of utility, we can now recover the unobserved demand shock for the reference variety k of good g in time t relative to the *reference variety* l of a *reference good* h in time t. Using (26) and (44), we have

$$\frac{b_{kgt}}{b_{lht}} = \left(\frac{s_{ht}}{s_{gt}}\right)^{\frac{1}{\gamma-1}} \frac{\tilde{P}_{ht}}{\tilde{P}_{gt}} \equiv \alpha_{gt} \tag{45}$$

where α_{gt} consists of variables that only depend on observed or estimated variables. And the price index for the aggregate importable composite good is proportional to the unobserved demand shock of the reference variety l of the reference good h and a psuedo-price index $\tilde{P}_{Mt}$ for the aggregate importable composite good that only depends on variables either observed or estimated:

$$P_{Mt} = b_{lht} \left[\sum_g \alpha_{gt} \left(\left(\sum_{v=1}^{V_g} \alpha_{vgt} p_{vgt}^{1-\sigma_g}\right)^{\frac{1}{1-\sigma_g}}\right)^{1-\gamma}\right]^{1/(1-\gamma)} \equiv b_{lht} \tilde{P}_{Mt}. \tag{46}$$

Moving to the upper level of utility, we can now recover the unobserved demand shock for the reference variety l of the reference good h as a function of expenditure shares and price indices for the aggregate importable and aggregate domestic composite goods. Using

(27) and (46), we have

$$b_{lht} = \left(\frac{s_{Dt}}{s_{Mt}}\right)^{\frac{1}{\kappa-1}} \frac{P_{Dt}}{\tilde{P}_{Mt}} \equiv \alpha_t \tag{47}$$

where α_{gt} consists of variables that only depend on observed or estimated variables. While we have focused on relative unobserved demand shocks, solving b_{lht} using (47) would then determine the level of the unobserved demand shock for the reference variety k of each good g using (45) and in turn the level of the unobserved demand shock for any variety of good g using (43). The problem is finding suitable data for the *level* of the price index for the domestic composite good in any given year.

Nevertheless, suppose we know there exists a good h and year y in which the unobserved demand shock for variety l is $b_{lht} \approx 1$ (i.e. a "zero" demand shock given b_{lht} enters utility multiplicatively). For example, perhaps we can manually identify a particular good in a particular year where the non-linear system of equations (41)-(42) is easy to solve. As just described, we could then use (45) and (43) to recover the level of the unobserved demand shocks of all varieties of all goods in year y. Moreover, we could also apply a two-step approach to recover the level of the domestic price index P_{Dt} in every year. First, manipulating (47) reveals

$$P_{Dy} \approx \left(\frac{s_{My}}{s_{Dy}}\right)^{\frac{1}{\kappa-1}} \tilde{P}_{My}. \tag{48}$$

Second, combining this price index level P_{Dy} in the reference year y with data such as a Producer Price Index from a national statistical agency would deliver price index levels P_{Dt} in all years $t \neq y$. Thus, we could then use (45) and (43) to recover the level of the unobserved demand shocks of all varieties of all goods in all years.

While we have just outlined two alternative procedures to recover the unobserved demand shocks, recovering all of these shocks reliably is not straightforward. Moreover, even if we could recover them reliably, one is still left with solving a high-dimensional and highly non-linear system of equations.

A.2 Solving the Model using Hat Algebra

Following Fajgelbaum et al. (2020), this section uses hat algebra (as opposed to *exact* hat algebra) to characterize the first order approximation of the equilibrum. Thus, the endogenous variables are log changes $\hat{x} \equiv d\ln x = \frac{1}{x}dx$. Specifically the model's endogenous variables are the proportional changes in (i) the equilibrium price indices $\hat{P}_t$, $\hat{P}_{Mt}$, and $\hat{P}_{gt}$, (ii) the equilibrium tariff-exclusive variety-level prices $\hat{p}^*_{vgt}$, and (iii) the equilibrium expenditure levels $\hat{E}_{Mt}$ and $\hat{E}_{gt}$. These endogenous variables depend on the following proportional changes in exogenous variables: (i) aggregate expenditure $\hat{E}_t$ (this could be endogenized to include tariff revenue) and the price index for the aggregate domestic composite good $\hat{P}_{Dt}$, (ii) legislated tariffs $\hat{\tau}_{vgt}$ and $\hat{f}_{vgt}$, and (iii) structural import demand and export supply shocks $\hat{b}_{vgt}$ and $\tilde{\eta}_{vgt}$.

A couple of modeling features deserve attention. First, we abstract from how tariffs affect domestic production and, in turn, interact with the importable side of consumer demand to affect the price index for the aggregate domestic composite good. Second, we abstract the impact of changes in tariff revenue on aggregate expenditure. Together, these two abstractions imply $\hat{E}_t$ and $\hat{P}_{Dt}$ are exogenous.

We begin deriving the hat form of the price indices. For the price index of good g at time t, differentiating (18) and using (25) reveals

$$\hat{P}_{gt} = \sum_{v=1}^{V_g} s_{vgt}\hat{p}_{vgt}. \tag{49}$$

For the price indices of the aggregate importable composite good and aggregate composite good, differentiating (23) and (24) while using (26) and (27) reveals

$$\hat{P}_{Mt} = \sum_{g=1}^{G} s_{gt}\hat{P}_{gt} \tag{50}$$

$$\hat{P}_t = s_{Mt}\hat{P}_{Mt} + s_{Dt}\hat{P}_{Dt}. \tag{51}$$

To find $\hat{p}^*_{vgt}$, we rely on the equilibrium tariff-exclusive price in equation (31). Since the equilibrium variety-level price depends on good-level expenditure E_{gt}, we first use the fact that $E_{gt} = E_{Mt}\left(\frac{P_{gt}}{P_{Mt}}\right)^{1-\gamma}$ and $E_{Mt} = E_t\left(\frac{P_{Mt}}{P_t}\right)^{1-\kappa}$ to see that

$$\hat{E}_{gt} = \hat{E}_t + (\kappa - 1)\,\hat{P}_t + (\gamma - \kappa)\,\hat{P}_{Mt} - (\gamma - 1)\,\hat{P}_{gt}. \tag{52}$$

Using (52) after taking the log and differentiating the equilibrium tariff-exclusive price in

(31), we have

$$\hat{p}^*_{vgt} = -\frac{\sigma_g\omega_g}{1+\sigma_g\omega_g}\frac{1}{1+AVE_{vgt}}dAVE_{vgt} + \frac{\omega_g}{1+\sigma_g\omega_g}\hat{b}_{vgt} + \frac{1}{1+\sigma_g\omega_g}\tilde{\eta}_{vgt}$$
$$+ \frac{\omega_g}{1+\sigma_g\omega_g}\left[\hat{E}_t + (\kappa-1)\hat{P}_t + (\gamma-\kappa)\hat{P}_{Mt} + (\sigma_g-\gamma)\hat{P}_{gt}\right] \tag{53}$$

where $\tilde{\eta}_{vgt} = d\eta_{vgt}$. Unlike when only considering advalorem tariffs, $AVE_{vgt} = \tau_{gt} + \frac{f_{gt}}{p^*_{vgt}}$ is endogenous in the presence of specific tariffs because it depends on p^*_{vgt}. Nevertheless, we can substitute $AVE_{vgt} = \tau_{gt} + \frac{f_{gt}}{p^*_{vgt}}$ in (53) to yield

$$\hat{p}^*_{vgt} = -\frac{\sigma_g\omega_g\varphi^{-1}_{vgt}}{1+\sigma_g\omega_g}\widetilde{AVE}_{vgt} + \frac{\omega_g\varphi^{-1}_{vgt}}{1+\sigma_g\omega_g}\hat{b}_{vgt} + \frac{\varphi^{-1}_{vgt}}{1+\sigma_g\omega_g}\tilde{\eta}_{vgt}$$
$$+ \frac{\omega_g\varphi^{-1}_{vgt}}{1+\sigma_g\omega_g}\left[\hat{E}_t + (\kappa-1)\hat{P}_t + (\gamma-\kappa)\hat{P}_{Mt} + (\sigma_g-\gamma)\hat{P}_{gt}\right] \tag{54}$$

where $\widetilde{AVE}_{vgt}$ and φ^{-1}_{vgt} were defined in (34) and (38). Note that $\varphi^{-1}_{vgt} = 1$ in the special case of $f_{vgt} = 0$ but $\varphi^{-1}_{vgt} < 1$ when $f_{vgt} > 0$.

We now have the simultaneous equation system of endogenous variables $\left\{\hat{P}_t, \hat{P}_{Mt}, \hat{P}_{gt}, \hat{p}^*_{vgt}\right\}$ and exoegnous variables $\left\{\widetilde{AVE}_{vgt}, \hat{b}_{vgt}, \tilde{\eta}_{vgt}, \hat{P}_{Dt}, \hat{E}_t\right\}$ governed by equations (49), (50), (51), and (54). Note that there are G price indices and associated equations for $\hat{P}_g$ for $g = 1, ..., G$ and $V = \sum_{g=1}^{G} V_g$ prices and associated equations for $\hat{p}_{vgt}$ for $v = 1, ..., V_g$ and $g = 1, ..., G$.

The system of equations characterizing the equilibrium is:

$$\begin{bmatrix} \hat{P}_t \\ \hat{P}_{Mt} \\ \left[\hat{P}_g\right]_{G\times 1} \\ \left[\hat{p}^*_{v,g,t}\right]_{V\times 1} \end{bmatrix} = [A_{yt} \mid A_{xt}] \begin{bmatrix} \hat{P}_t \\ \hat{P}_{Mt} \\ \left[\hat{P}_g\right]_{G\times 1} \\ \left[\hat{p}^*_{vgt}\right]_{V\times 1} \\ \left[\widetilde{AVE}_{vgt}\right]_{V\times 1} \\ \left[\hat{b}_{vgt}\right]_{V\times 1} \\ \left[\tilde{\eta}_{vgt}\right]_{V\times 1} \\ \hat{P}_{Dt} \\ \hat{E}_t \end{bmatrix} \tag{55}$$

where

$$A_{yt} = \begin{bmatrix} 0 & s_{Mt} & 0_{1\times G} & 0_{1\times V} \\ 0 & 0 & [s_{gt}]_{1\times G} & 0_{1\times V} \\ 0_{G\times 1} & 0_{G\times 1} & 0_{G\times G} & [\beta_{0vgt} s_{vgt}]_{G\times V} \\ [\beta_{1vgt}]_{V\times 1} & [\beta_{2vgt}]_{V\times 1} & [\beta_{3vgt}]_{V\times G} & 0_{V\times V} \end{bmatrix}$$

$$A_{xt} = \begin{bmatrix} 0_{1\times V} & 0_{1\times V} & 0_{1\times V} & s_{Dt} & 0 \\ 0_{1\times V} & 0_{1\times V} & 0_{1\times V} & 0 & 0 \\ [s_{vgt}]_{G\times V} & 0_{G\times V} & 0_{G\times V} & 0_{G\times 1} & 0_{G\times 1} \\ [\beta_{4vgt}]_{V\times V} & [\beta_{5vgt}]_{V\times V} & [\beta_{6vgt}]_{V\times V} & 0_{V\times 1} & [\beta_{5vgt}]_{V\times 1} \end{bmatrix}$$

and $\beta_{0vgt} = \left(1 - STS_{vgt}\frac{AVE_{vgt}}{1+AVE_{vgt}}\right)$, $\beta_{1vgt} = (\kappa - 1)\,\psi_{vgt}$, $\beta_{2vgt} = (\gamma - \kappa)\,\psi_{vgt}$, $\beta_{3vgt} = (\sigma_g - \gamma)\,\psi_{vgt}$, $\beta_{4vgt} = -\sigma_g \psi_{vgt}$, $\beta_{5vgt} = \psi_{vgt}$, $\beta_{6vgt} = \frac{1}{\omega_g}\psi_{vgt}$, and $\psi_{vgt} = \frac{\omega_g \varphi_{vgt}^{-1}}{1+\sigma_g \omega_g}$.

Letting $A_t = [A_{yt} \,\vdots\, A_{xt}]$, $y_t = \begin{bmatrix} \hat{P}_t & \hat{P}_{Mt} & \left[\hat{P}_g\right]_{1\times G} & \left[\hat{p}^*_{v,g,t}\right]_{1\times V} \end{bmatrix}'$, and $x_t = \begin{bmatrix} \left[\widetilde{AVE}_{vgt}\right]_{1\times V} & \left[\hat{b}_{vgt}\right]_{1\times V} & [\tilde{\eta}_{vgt}]_{1\times V} & \hat{P}_{Dt} & \hat{E}_t \end{bmatrix}'$, we can simply write the system of equations (55) as

$$\begin{aligned} y_t &= A_t \begin{bmatrix} y_t \\ x_t \end{bmatrix} \\ y_t &= (I - A_{yt})^{-1} A_{xt} x_t \end{aligned} \tag{56}$$

which defines the the endogenous variables $y_t = \begin{bmatrix} \hat{P}_t & \hat{P}_{Mt} & \left[\hat{P}_g\right]_{1\times G} & \left[\hat{p}^*_{v,g,t}\right]_{1\times V} \end{bmatrix}'$ in terms of the exogenous variables $x_t = \begin{bmatrix} \left[\widetilde{AVE}_{vgt}\right]_{1\times V} & \left[\hat{b}_{vgt}\right]_{1\times V} & [\tilde{\eta}_{vgt}]_{1\times V} & \hat{P}_{Dt} & \hat{E}_t \end{bmatrix}'$ and the parameters in A_t. Even though the system (56) uses very large matrices with our data and requires matrix inversion, MATLAB solves the system with relatively little computational burden.

A.3 Solving the Model using Exact Hat Algebra

To carry out our exact hat algebra analysis, we redefine our hat notation so that $\hat{z} \equiv \frac{z'}{z}$ where z' is the counterfactual value and z is the initial value of a variable z. We will characterize the endogenous variables $\hat{y}$ as a function of the exogenous variables $\hat{x}$.

We start by deriving the exact hat form of the price indices. Using the well-known results in the literature due to Sato (1976) and Vartia (1976), we have

$$\hat{P}_t = \prod_{k=M,D} \hat{P}_{kt}^{w_{kt}} \tag{57}$$

$$\hat{P}_{Mt} = \prod_{g \in G} \hat{P}_{gt}^{w_{gt}} \tag{58}$$

$$\hat{P}_{gt} = \prod_{v \in V_g} \hat{p}_{vgt}^{w_{vgt}}. \tag{59}$$

where the exact weights are

$$w_{vgt} = \frac{s_{vgt}\left(\hat{s}_{vgt} - 1\right)/\ln \hat{s}_{vgt}}{\sum_g s_{vgt}\left(\hat{s}_{vgt} - 1\right)/\ln \hat{s}_{vgt}} \tag{60}$$

$$w_{gt} = \frac{s_{gt}\left(\hat{s}_{gt} - 1\right)/\ln \hat{s}_{gt}}{\sum_g s_{gt}\left(\hat{s}_{gt} - 1\right)/\ln \hat{s}_{gt}} \tag{61}$$

$$w_{Mt} = \frac{s_{Mt}\left(\hat{s}_{Mt} - 1\right)/\ln \hat{s}_{Mt}}{\sum_{k=M,D} s_{kt}\left(\hat{s}_{kt} - 1\right)/\ln \hat{s}_{kt}} \tag{62}$$

and the exact hat tariff-inclusive expenditure shares are

$$\hat{s}_{vgt} = \frac{\left(\widehat{1 + AVE_{vgt}}\right)^{1-\sigma} \hat{p}_{vgt}^{*1-\sigma}}{\sum_{v=1}^{V_g} s_{vgt}\left(\widehat{1 + AVE_{vgt}}\right)^{1-\sigma} \hat{p}_{vgt}^{*1-\sigma}} \tag{63}$$

$$\hat{s}_{gt} = \frac{\hat{P}_{gt}^{1-\sigma}}{\sum_{g=1}^{G} s_{gt}\hat{P}_{gt}^{1-\sigma}} \tag{64}$$

$$\hat{s}_{Mt} = \frac{\hat{P}_{Mt}^{1-\sigma}}{\sum_{k=M,D} s_{kt}\hat{P}_{kt}^{1-\sigma}} \tag{65}$$

Using (19), the exact hat form for variety-level tariff-inclusive prices is

$$\hat{p}_{vgt} = \hat{p}_{vgt}^{*}(\widehat{1 + AVE_{vgt}}). \tag{66}$$

For the exact hat form of tariff-exclusive variety-level prices, we first derive an exact hat form for good-level equilibrium expediture E_{gt}. Using the fact that $E_{gt} = E_{Mt}\left(\frac{P_{gt}}{P_{Mt}}\right)^{1-\gamma}$ and

$E_{Mt} = E_t \left(\frac{P_{Mt}}{P_t}\right)^{1-\kappa}$, observed expenditure E_{gt} is

$$E_{gt} = E_t P_t^{\kappa-1} P_{Mt}^{\gamma-\kappa} P_{gt}^{1-\gamma} \tag{67}$$

and in exact hat form

$$\hat{E}_{gt} = \hat{E}_t \hat{P}_t^{\kappa-1} \hat{P}_{Mt}^{\gamma-\kappa} \hat{P}_{gt}^{1-\gamma}. \tag{68}$$

Imposing $b'_{vgt} = b_{vgt}$ and $\eta'_{vgt} = \eta_{vgt}$ and using (31), the exact hat form for variety-level tariff-exclusive prices is

$$\hat{p}^*_{vgt} = (\widehat{1 + AVE}_{vgt})^{-\frac{\sigma_g \omega_g}{1+\sigma_g \omega_g}} \left(\hat{E}_t \hat{P}_t^{\kappa-1} \hat{P}_{Mt}^{\gamma-\kappa} \hat{P}_{gt}^{\sigma_g - \gamma}\right)^{\frac{\omega_g}{1+\sigma_g \omega_g}}. \tag{69}$$

Importantly, as noted in Appendix A.2, AVE_{vgt} depends on the endogenous variable p^*_{vgt}. In turn,

$$(\widehat{1 + AVE}_{vgt}) = (\widehat{1 + \tau_{gt}}) \left(1 - STS_{vgt} \frac{AVE_{vgt}}{1 + AVE_{vgt}}\right) + \frac{\hat{f}_{gt}}{\hat{p}^*_{vgt}} STS_{vgt} \frac{AVE_{vgt}}{1 + AVE_{vgt}}. \tag{70}$$

We now have an exact hat algebra system of equations that characterizes the counterfactual equilibrium. The endogenous variables are (i) the price indices $\hat{P}_t$, $\hat{P}_{Mt}$, $\left[\hat{P}_{gt}\right]_{G\times 1}$ and, using (66), the tariff-exclusive prices $\left[\hat{p}^*_{vgt}\right]_{V\times 1}$ characterized by (57)-(59) and (69), (ii) using (60)-(62), the expenditure shares $\hat{s}_{Mt}$, $[\hat{s}_{gt}]_{G\times 1}$, and $[\hat{s}_{vgt}]_{V\times 1}$ characterized by equations (63)-(65), and (iii) the AVEs $\left[(\widehat{1 + AVE}_{vgt})\right]_{V\times 1}$ characterized by equation (70). These endogenous variables depend on the following exogenous variables: (i) initial expenditure shares s_{Mt}, s_{Dt}, $[s_{gt}]_{G\times 1}$, $[s_{vgt}]_{V\times 1}$, (ii) initial values $[AVE_{vgt}]_{V\times 1}$ and $[STS_{vgt}]_{V\times 1}$, (iii) changes in legislated tariffs $\left[(\widehat{1 + \tau_{gt}})\right]_{V\times 1}$ and $\left[\hat{f}_{gt}\right]_{V\times 1}$, and (iv) changes in aggregate expenditure and the domestic price index $\hat{E}_t$ and $\hat{P}_{Dt}$.

The exact hat algebra system is substantially more complicated to solve than the hat algebra system. Not only is it highly non-linear compared to the linear hat algebra system, it also has roughly twice as many endogenous variables as the hat algebra system which was already a high-dimensional system. This is because the exact hat algebra system not only has the endogenous prices $\hat{P}_{gt}$ and $\hat{p}^*_{vgt}$ but also endogenous shares $\hat{s}_{gt}$ and $\hat{s}_{vgt}$ whereas the hat algebra system only has endogenous prices $\hat{P}_{gt}$ and $\hat{p}^*_{vgt}$. While the exact hat algebra change in price indices and prices require knowing the counterfactual change in expenditure shares, the hat algebra (i.e. first order approximation) of changes in price indices and prices only require knowing the initial expenditure shares. Moreover, the presence of specific tariffs add an additional layer of non-linearity to the exact hat algebra system. This can be seen from (69)-(70) which imply $\hat{p}^{*}_{vgt}$ enters the right hand side of (69) only in the presence of specific tariffs and does so non-linearly.

B Heteroskedastic elasticity estimator

We begin by double differencing the import demand and export supply curves – equations (28) and (30) – with respect to time and a reference variety k. The first difference eliminates time-invariant tastes for varieties and the second difference removes good level price indices. For the import demand curve, we thus have

$$\Delta^k \ln s^*_{vgt} = -(\sigma_g - 1)\Delta^k \ln p^*_{vgt} - \sigma_g \Delta^k \ln(1 + AVE_{vgt}) + \varepsilon^k_{vgt} \tag{71}$$

where

$$\begin{aligned}
\varepsilon^k_{vgt} &= (\varepsilon_{vgt} - \varepsilon_{vgt-1}) - (\varepsilon_{kgt} - \varepsilon_{kgt-1}) \\
&= ((\sigma_g - 1) P_{gt} + \ln b_{vgt} - (\sigma_g - 1) P_{gt-1} - \ln b_{vgt-1}) \\
&\quad - ((\sigma_g - 1) P_{gt} + \ln b_{kgt} - (\sigma_g - 1) P_{gt-1} - \ln b_{kgt-1}) \\
&= (\ln b_{vgt} - \ln b_{vgt-1}) - (\ln b_{kgt} - \ln b_{kgt-1}) \\
\varepsilon^k_{vgt} &= \Delta^k \ln b_{vgt}.
\end{aligned} \tag{72}$$

For the export supply curve we have

$$\Delta^k \ln p^*_{vgt} = \frac{\omega_g}{1 + \omega_g} \Delta^k \ln s^*_{vgt} + \delta^k_{vgt} \tag{73}$$

where

$$\begin{aligned}
\delta^k_{vgt} &= (\delta_{vgt} - \delta_{vgt-1}) - (\delta_{kgt} - \delta_{kgt-1}) \\
&= \frac{1}{1 + \omega_g} (\omega_g \ln E_{gt} + \eta_{vgt} - \omega_g \ln E_{gt-1} - \eta_{vgt-1}) \\
&\quad - \frac{1}{1 + \omega_g} (\omega_g \ln E_{gt} + \eta_{kgt} - \omega_g \ln E_{gt-1} - \eta_{kgt-1}) \\
&= (\ln \eta_{vgt} - \ln \eta_{vgt-1}) - (\ln \eta_{kgt} - \ln \eta_{kgt-1}) \\
\delta^k_{vgt} &= \Delta^k \ln \eta_{vgt}.
\end{aligned} \tag{74}$$

This system looks nearly identical to the standard F/BW system. The key differences are that our endogenous variables are measured in tariff-exclusive rather than tariff-inclusive prices and we have the extra AVE term in the demand equation.

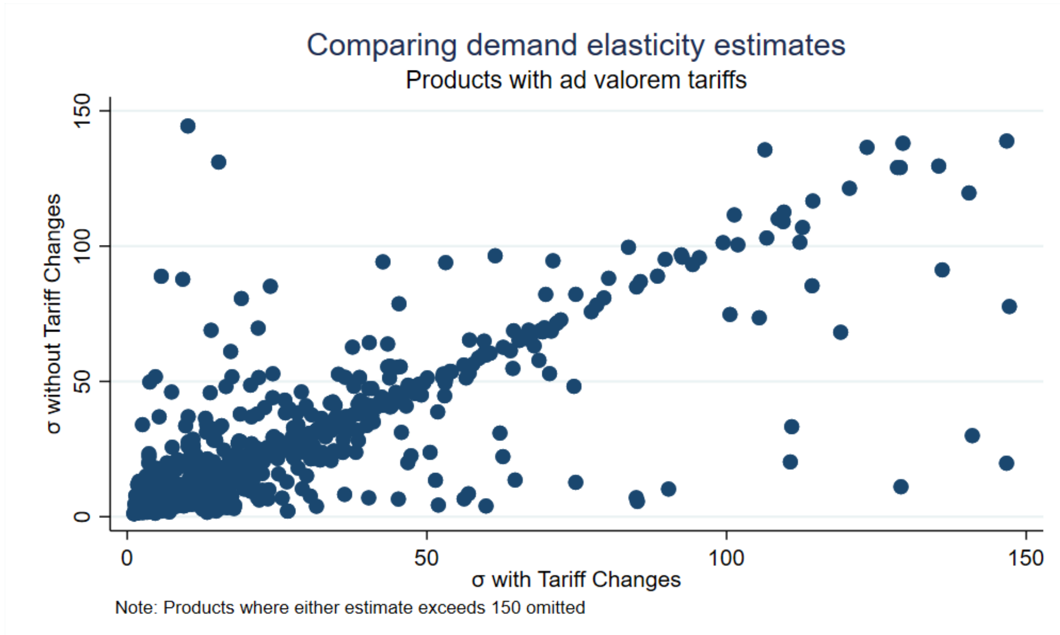
Multiplying (71) and (73) yields

$$\begin{aligned}\left(\Delta^k \ln p^*_{vgt}\right)^2 =& \theta_{g1}\left(\Delta^k \ln s^*_{vgt}\right)^2 + \theta_{g2}\Delta^k \ln p^*_{vgt}\Delta^k \ln s^*_{vgt} \\ &+ \theta_{g3}\Delta^k \ln p^*_{vgt}\Delta^k \ln\left(1+AVE_{vgt}\right) + \theta_{g4}\Delta^k \ln s^*_{vgt}\Delta^k \ln\left(1+AVE_{vgt}\right) + \frac{\varepsilon^k_{vgt}\delta^k_{vgt}}{\sigma_g - 1}\end{aligned} \tag{75}$$

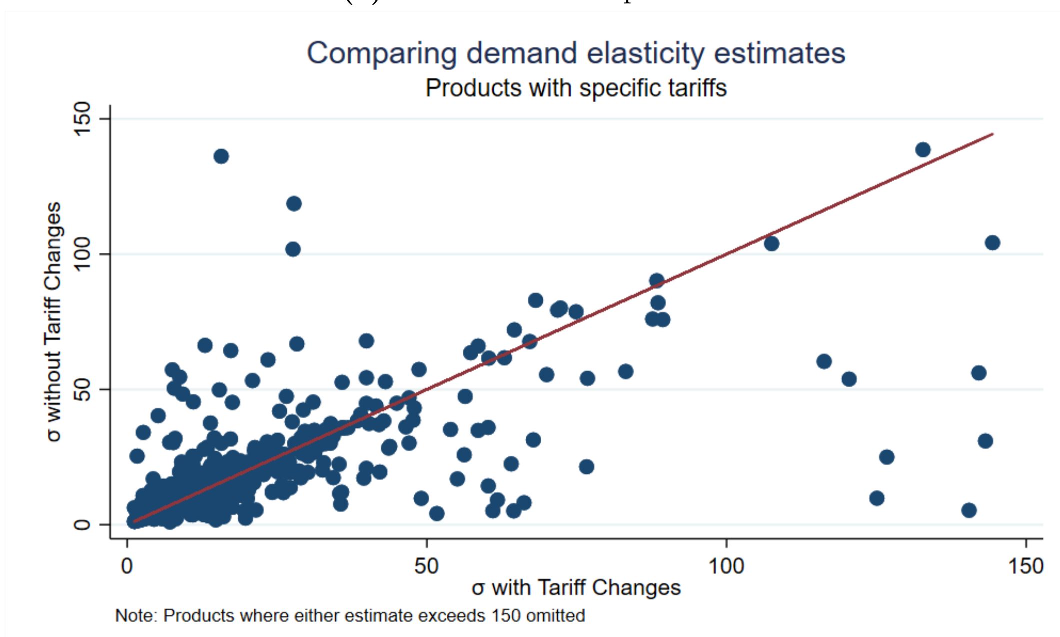
where $\theta_{g1} = \frac{\omega_g}{(1+\omega_g)(\sigma_g-1)}$, $\theta_{g2} = -\frac{\omega_g(2-\sigma_g)+1}{(1+\omega_g)(\sigma_g-1)}$, $\theta_{g3} = \frac{\sigma_g}{\sigma_g-1}$, and $\theta_{g4} = \frac{\sigma_g\omega_g}{(1+\omega_g)(\sigma_g-1)}$. While θ_{g1} and θ_{g2} take exactly the same functional form as the standard F/BW setup, we have the additional coefficients θ_{g3} and θ_{g4}. We solve (75) using our full sample from 1972-1988 using non-linear least squares and imposing $\omega_g > 0$ and $\sigma_g > 1$.

Figure 10 plots our estimate of the standard F/BW elasticities on the y-axis – denoted "σ without tariff changes" – against the our generalized elasticity estimates – denoted "σ with tariff changes". Both sets of elasticities are obtained from estimating (75) but the standard F/BW estimates impose $\theta_{g3} = \theta_{g4} = 0$ while our generalized elasticity estimates do not impose $\theta_{g3} = \theta_{g4} = 0$.

Overall, Figure 10 illustrates a modest empirical bias when ignoring specific and ad valorem tariffs in estimating the structural elasticity σ_g. The clustering along the 45-degree line illustrates the clear majority of products have very similar elasticity estimates regardless of whether one imposes $\theta_{g3} = \theta_{g4} = 0$. However, the bias is non-trivial for a considerable number of products, with a larger bias among varieties facing specific tariffs. Specifically, the respective mean absolute difference for the two elasticities is 16.6% and 10.9% for the respective sub-samples of products that only have ad valorem tariffs and have specific tariffs. The larger bias for specific tariff products is expected, as variation in the foreign price across varieties is enough to ensure $\Delta^k \ln\left(1+AVE_{jt}\right)$ is non-zero. Nevertheless, the bias for products facing only ad valorem tariffs suggests a role for preferential tariff margins across exporters that vary over time at the tariff-line level – e.g., for reasons related to the Generalized System of Preferences.



(a) Ad valorem tariff products



(b) Specific tariff products

Figure 10: Comparing demand Elasticity estimates

Notes: Figure plots estimates of σ estimated using equations (75) when imposing $\ln(1 + AVE_{jt}) = 0$ (y-axis) and when using $\ln(1 + AVE_{jt})$ from the data (x-axis). Products where either elasticity estimate exceeds 150 excluded from figure.