AVE Tariffs are a Misleading Measure of Tariff Policy*

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Abstract

We show that changes in ad valorem equivalent (AVE) tariffs are insufficient to sign the effect of tariffs on economic outcomes. Using a novel database of 1972-1988 US tariffs, we decompose aggregate AVEs into statutory and endogenous components. Pre-1979, inflation creates an "accidental liberalization" by eroding specific tariffs, while import composition changes mask 1980s Tokyo Round tariff cuts. Overall, as much as 70% of the variation in aggregate AVE changes is driven by mechanisms other than statutory tariff changes. Embedding specific tariffs in a hat algebra framework, we show that the accidental liberalization accounts for half of 1972-1988 AVE changes. The resulting impact on imports and price indices depends on whether price movements were driven by import demand or export supply shocks. If driven by export supply shocks, the accidental liberalization is associated with falling imports and rising import price indices.

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1 Introduction

The ad valorem equivalent (AVE) tariff, defined as the ratio of duties collected to the value of imports, has long been the most widely used measure of tariff protection (Anderson and Neary, 2005). Over the past decade, rising protectionism has led to a resurgence of interest in the effect of tariffs, both among trade economists (Amiti et al., 2019; Fajgelbaum et al., 2020; Flaaen and Pierce, 2024; Handley et al., 2025) and those concerned with the macroeconomic consequences of tariffs more broadly (Barnichon and Singh, 2025; Schmitt-Grohe and Uribe, 2025; den Besten and Känzig, 2026). This work has often placed particular emphasis on AVE tariffs. However, the AVE has a number of shortcomings that complicate its use as a standalone measure of tariff policy. In particular, endogenous changes in price levels and import shares affect the AVE in ways that produce an ambiguous relationship with domestic prices and imports. Put simply, the AVE tariff is an endogenous equilibrium object, and attempts to use it as a policy primitive risk misstating the effects of tariffs on price indices and imports—not only in magnitude but in direction.

In this paper, we demonstrate this point using a newly constructed database of US statutory tariffs and import values between 1972 and 1989, years covering the negotiation and implementation of the Tokyo Round of the General Agreement on Tariffs and Trade (GATT). The aggregate AVE tariff fell by roughly 2 percentage points in this period—a larger decline than those experienced under the Canada-US Free Trade Agreement (CUSFTA), the North American Free Trade Agreement (NAFTA), or the Uruguay Round. Further, the era witnessed substantial changes in price levels and import patterns, making it an ideal period in which to explore the determinants of the AVE tariff and their respective effects on economic outcomes of interest.

With these data we highlight the fact that relatively little of the variation in the aggregate AVE stems from statutory changes in ad valorem tariffs—as is often implicitly assumed when using AVEs. We find that the shifting composition of imports accounts for more than one-third of the variation in aggregate AVE changes in these years. Moreover, rising import prices in the presence of specific tariffs account for roughly as much of the decline in the AVE as statutory tariff changes. This non-policy variation cannot be treated as measurement error, as these components of tariff changes are endogenously related to outcomes of interest. Further, the relationship between changing AVEs and economic outcomes of interest will depend on the source of the AVE changes. Price increases driven by rising domestic demand, for instance, correspond to falling AVE and rising imports, while those driven by negative foreign supply shocks lead to reductions in both AVE and imports. Overall, as much as 70% of the variation in aggregate AVE changes in this period is driven by mechanisms other than statutory tariff changes, with ex ante ambiguous effects on imports and price indices. Our findings serve as a word of caution to researchers exploring the effects of tariffs on economic outcomes: assessing the relationship between AVE, imports, and prices requires that one know the source of variation in the AVE.

Our analysis begins with an apparent puzzle in the aggregate US AVE between 1972 and 1989. The Tokyo Round concluded in 1979, with negotiated tariff cuts phased in over the following eight years. This represented the first wholesale change in the US tariff code since the tariff reductions mandated by the Kennedy Round of the GATT finished in 1972.[1] One would thus expect relative stability in the AVE during the 1970s and a subsequent decline during the 1980s. However, Figure 1 shows that the *opposite* is true. That is, despite the absence of meaningful tariff policy changes, the US AVE fell sharply throughout the 1970s and, despite the successful conclusion of the Tokyo Round negotiations, the AVE *rose* slightly during the 1980s.

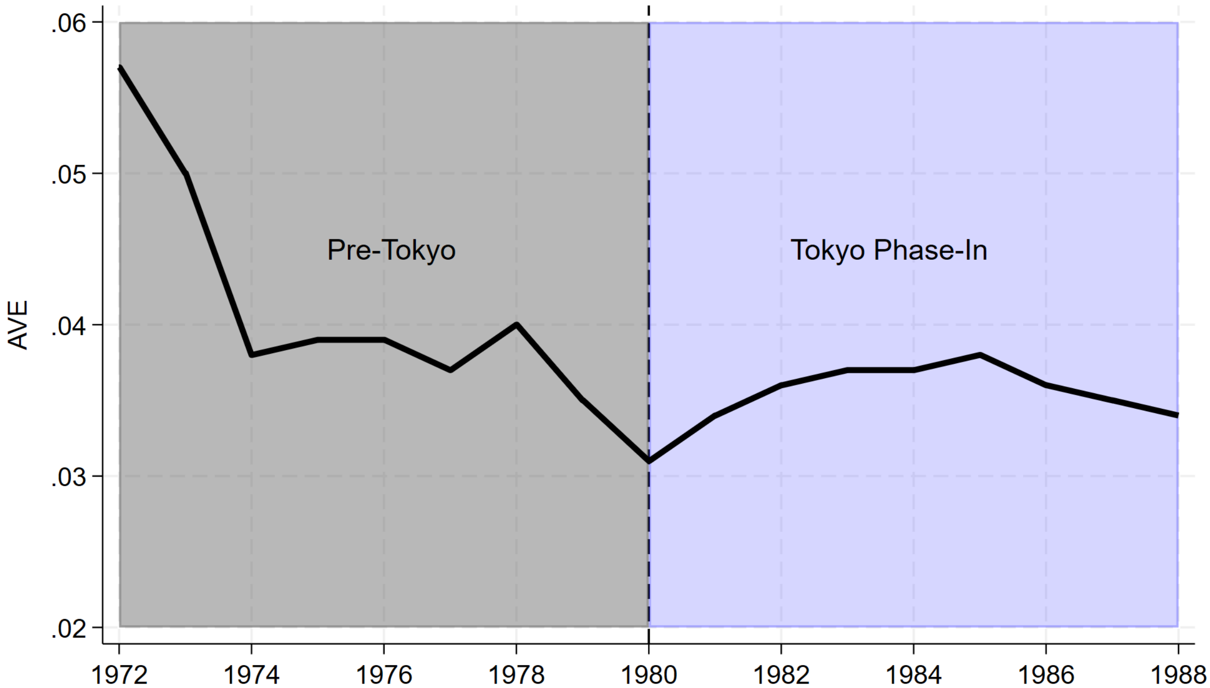
To reconcile these changes with the common understanding of trade policy in this period and to highlight how different sources of AVE changes can confound the analysis of tariff policy, we construct a database of US tariffs for all imported products between 1972 and 1988. The database details statutory tariffs—distinguishing between ad valorem, specific, and compound tariffs—rather than relying on the product-level AVEs used extensively throughout the literature.[2] With these data in hand, we decompose changes in the aggregate AVE into changes in tariff-line AVEs, the composition of imports, and the set of imported products. This decomposition reveals two distinct liberalizations during this era. First, the statutory liberalization dictated by the Tokyo Round reduces US tariffs by nearly 25% between 1980 and 1987, largely along lines prescribed by the so-called "Swiss Formula".[3] However, en-

[1]The "Nixon Shock" altered statutory US tariffs after the Kennedy Round negotiations concluded, but lasted for only four months in 1971, prior to the start of our sample (Irwin, 2012).

[2]These data encompass not only the Tokyo Round liberalization, but also provide column 2 "non-normal trade relations" tariffs, temporary legislative tariff changes, and the introduction of preferential tariff programs. We combine these data with data on trade flows from the Census Annual Import Data Bank (IDB) over the same period and correct several important inaccuracies in the IDB. An additional contribution of this paper is to make this joint import and tariff dataset available to researchers.

[3]The Tokyo Round was the first use of the Swiss Formula for tariff reductions. The formula, proposed by the Swiss delegation to the GATT in 1976, defines a final tariff $Z = \frac{AX}{A+X}$, 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 for the US, though the formula was not followed for all products. For more details, see Greenland et al. (2026) and https://www.wto.org/english/tratop_e/agric_e/agnegs_swissformula_e.htm.

Figure 1: Aggregate US Ad Valorem Equivalent Tariff



Notes: US ad valorem equivalent tariff (AVE), defined as duties collected divided by total imports. Data from USITC: https://www.usitc.gov/documents/dataweb/ave_table_1891_2016.pdf.

dogenous changes in the composition of imports that accompany the tariff cuts mask this liberalization at the aggregate level. Second, we document a substantial "accidental" liberalization in the years preceding the Tokyo Round driven by changes in tariff-exclusive import prices. Rapidly rising prices in the presence of specific tariffs, which cover one-third of US tariff lines at the beginning of our sample, reduce the AVE on such tariff lines from 7% to 4%.[4]

The decomposition highlights two main issues relevant for trade policy studies. First, when analysis takes place at a more aggregate level than the tariff line, AVE changes reflect not only changes in statutory tariffs but also endogenous, non-policy changes in the composition of imports. Although microeconomic studies commonly fix the composition of imports as of some initial period, the literature focusing on more aggregate outcomes often

[4] These price changes can be further separated into foreign-currency-denominated price shocks and exchange rate movements. While we focus on the combined effect in the main text, Appendix F extends our framework to show that a similar effect would have persisted throughout the 1980s in the absence of the dramatic 1980s movements in the US dollar exchange rate.

does not (Estevadeordal et al., 2003; Clemens and Williamson, 2004; DeJong and Ripoll, 2006; Giuliano et al., 2013; Furceri et al., 2022; Schmitt-Grohe and Uribe, 2025; Barnichon and Singh, 2025; den Besten and Känzig, 2026).[5] Additionally, analyses relying on "off-the-shelf" tariff data that has been aggregated to the product or industry level suffer from this issue as well. Perhaps most notably, widely used tariff data from the World Integrated Trade Solution (WITS) defaults to providing data at the six-digit Harmonized System (HS) level, rather the tariff line-level of each country, which is typically more disaggregate. Second, the presence of specific tariffs implies that AVE changes are not sufficient for identifying import effects even at the tariff-line level. The intuition is straightforward: higher tariff-exclusive prices reduce the AVE of specific tariffs but also increase tariff-inclusive import prices and hence can reduce imports.

While the distinction between policy and non-policy variation in tariff levels has been emphasized previously (Van Cott and Wipf, 1983; Crucini, 1994; Irwin, 1998a,b; Greenland and Lopresti, 2024; Klein and Meissner, 2024), analysis quantifying the aggregate importance of this distinction for imports and prices has been sparse or ad hoc.[6] A likely explanation for this gap in the literature is that the additive nature of specific tariffs complicates quantitative analysis by clashing with the multiplicative structure of standard general equilibrium models.[7] To address this, we generalize the CES demand and Armington production structure adopted by Fajgelbaum et al. (2020) and demonstrate that one can employ modern "hat" and "exact hat" algebra techniques to conduct counterfactual analysis in the presence of specific tariffs. Our counterfactuals further underscore the importance of accounting for the source of AVE variation: approximately half of the change in tariff-inclusive import prices from falling tariff-line AVEs between 1972 and 1988 is driven by inflation rather than statutory tariff changes.[8,9] Given our estimated elasticities, our model interprets the higher

[5]Given their aggregate data, Clemens and Williamson (2004) acknowledge issues related to changing prices but they can neither address this nor issues driven by changing import composition.

[6]Policymakers have long recognized the importance of specific relative to ad valorem tariffs. Indeed, the USITC has published multiple studies exploring the feasibility of converting the tariff schedule entirely to ad valorem tariffs. See `https://www.usitc.gov/publications/332/pub896.pdf` and `https://www.usitc.gov/publications/docs/pubs/332/pub3318/Simprpt.PDF`.

[7]Moreover, empirical researchers typically omit specific tariffs, convert them to AVEs, or rely on data sources that do this implicitly (Topalova, 2010; Head and Mayer, 2014; Caliendo and Parro, 2015; Hakobyan and McLaren, 2016; Pierce and Schott, 2016; Dix-Carneiro and Kovak, 2017; Handley and Limão, 2017; Greenland et al., 2019; Besedes et al., 2020; Alessandria et al., 2025).

[8]In subsequent work, Geschonke (2024) implements our decompositions on an annual basis to evaluate the reduced-form relationship between AVE tariffs and economic growth in Germany over the 1880-1910 period. Also in subsequent work, Phillips (2024) analyzes the impacts of a passive currency devaluation on the AVE of specific tariffs in the 2000s.

[9]In our model, the change in welfare is proportional to the change in the importable price index. However,

observed prices that drive this "accidental liberalization" as reflecting negative export supply shocks that ultimately *increase* prices and *reduce* imports.

While our primary focus is on the years surrounding the Tokyo Round, the issues accompanying the use of the AVE as a policy variable persist into later years. In an appendix, we show that compositional changes explain a majority of the variation in the aggregate US AVE between 1989 and 2016. Further, while specific tariffs have fallen as a share of tariff lines over time, they remain an important determinant of cross-sectional variation in AVEs: between one-third and one-half of the annual cross-sectional variance in tariff-line AVEs in these years is driven by the variance in specific tariff AVEs. This has important implications for standard empirical approaches that use cross-sectional variation in AVE levels as a source of identification.[10] As such, the issues we document in the context of the Tokyo Round will also confound analysis of trade policy in more modern settings.

Our work is related to the literature focused on data issues surrounding the estimation and quantification of trade policy. As emphasized by Goldberg and Pavcnik (2016), studies of trade policy often rely on detailed tariff-line level data. For the period since the advent of the Harmonized System in 1989, work by, e.g., Feenstra (1996), Feenstra et al. (2002), and Pierce and Schott (2012) provides US tariff data and concordances. Recent work by, e.g., Caliendo et al. (2023) and Teti (2024) has greatly improved available cross-country tariff data. We contribute to a group of studies focused on earlier years, in which average tariffs were considerably higher and more variable (Buzard et al., 2024; Greenland and Lopresti, 2024; Klein and Meissner, 2024; Acosta and Cox, 2025; Greenland et al., 2026; Madsen et al., 2025). Most similar to us in this regard, Acosta and Cox (2025) provide US tariff-line data for the years corresponding to the implementation of negotiated GATT MFN tariffs, including the Kennedy Round in 1972 and the Tokyo Round phase-ins between 1980 and 1987. We go beyond this by providing MFN tariffs as well as column 2 and preferential tariffs annually for the entire 1972-1988 period and synthesizing these tariff data with Census import data.

We also contribute to the numerous studies quantifying the effects of trade policy on US domestic economic outcomes. Studies of the 1989 Canada-US Free Trade Agreement (Trefler, 1993; Guadalupe and Wulf, 2010), the North American Free Trade Agreement (Caliendo and Parro, 2015; Hakobyan and McLaren, 2016; Choi et al., 2024), the US granting of Permanent

we do not model the production side of the economy in order to focus on the direct consumption-side effects of specific tariffs.

[10] Early examples of this approach come from Goldberg and Pavcnik (2005) and Topalova (2010). In the US, this strategy has been employed in the context of, e.g., NAFTA (Hakobyan and McLaren, 2016; Benguria, 2023; Choi et al., 2024) and PNTR (Pierce and Schott, 2016; Greenland et al., 2019).

Normal Trade Relations to China in 2001 (Pierce and Schott, 2016; Handley and Limão, 2017; Greenland et al., 2019), the 2001 Bush steel tariffs (Cox, 2025; Lake and Liu, 2025), and the Trump-era tariffs (Fajgelbaum et al., 2020; Flaaen and Pierce, 2024; Ignatenko et al., 2025) all rely on statutory tariff changes to establish causal links between trade policy and economic outcomes. In contrast, we emphasize the importance and implications of changing AVEs driven by both statutory tariff changes and the inflationary erosion of specific tariffs.

Additionally, we add to the literature focused on the role of specific tariffs in the US tariff code, the majority of which emphasizes the importance of specific tariffs in shaping aggregate tariff levels, import patterns, and labor market outcomes during the first half of the 20th century US (Crucini, 1994; Irwin, 1998a,b; Bond et al., 2013; Greenland and Lopresti, 2024).[11] Our analysis shows that specific tariffs are important for empirical and quantitative studies into the 21^{st} century, far more recently than previously demonstrated. We also emphasize that the presence of specific tariffs can affect the sign of the relationship between AVEs, prices, and imports. The predicted nature of these relationships will depend on assumptions about whether price-induced AVE changes are driven by export supply or import demand shocks, as well as the correlation structure between these shocks. While the present paper does not model the production side of the economy and hence wages, these assumptions are also needed for interpreting the relationships between AVEs, wages, and welfare. The presence of specific tariffs will thus also complicate the relationship between AVE and welfare.

Finally, our paper is related to a number of studies analyzing the effects of evolving US trade policy in the 1970s and 1980s. Among these, only Van Cott and Wipf (1983) address the role of inflation on AVEs. The authors provide back-of-the-envelope estimates of the effects of inflation on aggregate AVEs in the pre-Tokyo era, but do not assess the effect on imports or address statutory tariff changes. Deardorff and Stern (1979, 1981) provide prospective analyses of the effects of the Tokyo Round before its implementation, while Brown (1988) provides estimates of the effects of GSP implementation in a computable general equilibrium model. Trefler (1993) focuses on the effects of non-tariff barriers during the 1980s rather than tariff changes under the Tokyo Round. Relative to these studies, we provide a comprehensive analysis of the years surrounding the Tokyo Round through a structural model of the various sources of tariff changes in the era.

[11] Other studies focus more generally on per-unit trade costs. Hummels and Skiba (2004) and Feenstra and Romalis (2014) emphasize the importance of per-shipment trade costs in determining both trade volume and export quality. Irarrazabal et al. (2015) detail a procedure to estimate additive trade costs. Sørensen (2014) shows that the gains from trade are larger in the presence of such barriers.

The paper proceeds as follows. Section 2 describes our tariff database and provides a decomposition of tariff changes into their statutory and endogenous components at the aggregate and tariff-line levels. Section 3 embeds specific tariffs in a CES framework and describes counterfactual analysis in hat algebra and exact hat algebra settings. We also outline an approach to estimating elasticity parameters based on the Swiss Formula. Section 4 presents a number of counterfactual estimates to quantify the role of statutory and inflationary changes in AVEs. Section 5 concludes.

2 The Evolution of US Tariff Policy 1972-1988

2.1 A Novel Dataset of Statutory Tariffs

Our sample period begins in 1972 and ends in 1988. During this time, the global trading system operated under two distinct GATT regimes. Kennedy Round negotiations were completed in 1967, with negotiated tariff cuts fully phased in by 1972. Tokyo Round negotiations concluded in 1979, with tariff reductions implemented over a phase-in schedule that lasted until 1987. Statutory tariff schedules for the US in this period, including phase-ins, were covered by the Tariff Schedule of the United States Annotated (TSUSA).[12]

Currently, no dataset contains annual product-level US statutory tariffs for these years. Instead, work relying on tariffs in this era has largely taken an indirect approach based on the dataset constructed by Feenstra (1996), which provides annual US imports using data produced by the US Census in the Annual Import Data Bank (IDB). Product-level AVEs can be calculated in these data as duties collected divided by import values. However, this approach ignores the distinction between tariffs specified in percentage (ad valorem) terms and those specified in per-unit (specific) terms. As such, one cannot determine whether changing AVEs are driven by changes in statutory tariffs or by price movements in the presence of specific tariffs. This is a non-trivial concern in this era, as price levels doubled throughout the 1970s and specific tariffs were pervasive.

To address this, we construct a dataset that is more complete, cleaner, more disaggregate, and more consistent across time than any currently available for these years.[13] It is more complete for two reasons. First, it provides both statutory tariffs and imports. In particular,

[12] Unlike the subsequent Uruguay Round negotiations, Tokyo Round negotiations were over tariffs themselves, rather than over the upper bound on tariffs, known as "tariff bindings".

[13] We detail construction of the dataset in Appendix C.

it is the only dataset that provides statutory ad valorem tariffs, τ_{vgt}, and specific tariffs, f_{vgt}, by exporter v, five-digit TSUSA good g, and year t during our sample period. In doing so, it recognizes that such tariffs may differ by "column 1" or "column 2" status, or may fall under one of the special preferential tariff programs used by the US in this period: the Generalized System of Preferences (GSP), the Least Developed Developing Countries (LDDC) program, the Caribbean Basin Economic Recovery Act of 1983 (CBERA), or the Israel Free Trade Agreement (IFTA).[14] Second, our dataset provides statutory tariffs for 1972 and 1973, which cannot be done using Feenstra (1996) as 1972 and 1973 duties are not provided therein. We overcome this limitation by digitizing the 1972 TSUSA, which also dictates the 1973 tariff schedule.

Our dataset is cleaner than Feenstra (1996) for two reasons. First, he notes that approximately 20% of 1977 observations are missing from the IDB data used to construct his dataset. He addresses this issue by digitizing hard copies of the import data. However, researchers now have digital access to the archival version of the 1977 IDB that includes these previously missing observations. While the digitization in Feenstra (1996) is quite accurate, it is naturally imperfect. We overcome this by using the archival IDB data. Second, we provide the correct statutory tariffs in cases in which the IDB incorrectly calculates duties collected.[15] For example, a non-trivial subset of 1986 duties reported in the IDB use statutory tariffs from the 1985 TSUSA rather than the 1986 TSUSA.[16] Additionally, duties for a substantial number of observations with non-zero specific tariffs are calculated using a different unit of quantity than specified in the TSUSA.[17]

[14]Introduced in 1976, GSP provides duty-free access to a list of developing countries for certain products. Rather than waiting until 1987 for the fully phased-in Tokyo Round tariff cuts, the LDDC program gave a subset of developing countries access to the 1987 tariffs in 1980. The CBERA began providing tariff-free access to Caribbean countries on certain products in 1984. Finally, the IFTA began eliminating tariffs on imports from Israel in 1985. Appendix Figure A.1 shows these programs cover approximately 30-50% of tariff lines and eligible imports account for approximately 3-5% of US imports.

[15]Documentation accompanying the IDB clearly states that the calculated duties variable is a variable *calculated* by the Census, and that users should use the variable with caution. Notably, it is not a variable that is taken from customs forms.

[16]The IDB data include "rate provision codes", which describe the tariff program under which a particular product entered – column 1, column 2, GSP, etc. The system of rate provision codes changed in 1986, but the 1986 import data sometimes contains rate provision codes for both systems within a given five-digit TSUSA code. While observations using the new system use tariffs from the 1986 TSUSA, observations using the old system use 1985 tariffs. For example, consider TSUSA product 11218 "Salmon, not in oil, in airtight container". Consistent with the 1985 and 1986 TSUSA PDFs, dividing column 1 duties collected by import value yields ad valorem tariffs of 4.1% for observations with the old-system rate provision code and 3.6% for observations with the new-system rate provision code. This problem is pervasive, as 38% of 1986 dutiable imports use the old rate provision codes.

[17]Examples include the five-digit TSUSA oil-related products 47505, 47510, 47525, 47530, 47535, 47545,

Finally, our dataset is both disaggregate and consistent across time. The unit of observation in our data is a country-good-year-port rather than a country-good-year, and we report import quantities using a time-consistent unit of quantity.[18] While relatively uncommon, the IDB data contain instances in which the unit of quantity changes over time, yielding unit values that are incomparable. We correct such cases.[19]

Table 1 presents descriptive statistics for tariffs and imports from our database separately by one-digit TSUSA sector as of 1979. The share of tariff lines in each sector typically varies between 15% and 21%, with the exceptions of "Wood & Paper" (6%) and "Non-Metal Minerals & Products" (7%). The distribution of imports across sectors is somewhat different, with a notably smaller share of imports accounted for by "Textiles" (5%), "Miscellaneous Manufactured Products" (10%) and "Agriculture" (12%) and a larger share accounted for by "Metals & Metal Products" (49%). There is substantial heterogeneity in tariff prevalence across sectors. Most sectors have duty-free access on 15-25% of tariff lines, with the notable exception of "Textiles" (5%) and "Miscellaneous Manufactured Products" (5%), for which a much higher share of lines face tariffs. Moreover, these two sectors exhibit the highest AVE tariff on dutiable goods (13-16%), whereas "Wood & Paper" and "Chemicals", with the lowest AVEs (5-6%), report the greatest share of duty-free lines.

Specific tariffs are pervasive in this period: more than 60% of tariff lines in "Agriculture" have specific tariffs, and the share is above 20% in all sectors. As we discuss below in detail, the high share of specific tariffs in agriculture makes it especially prone to the inflationary erosion of AVE; indeed, its mean AVE falls by 3.2 percentage points between 1972 and 1979. In contrast, given the nature of the Swiss Formula, the largest declines in post-Tokyo AVE come in the two sectors with the highest initial AVE: "Textiles", and "Miscellaneous Manufactured Products" with tariff reductions of 6.4 and 6.1 percentage points, respectively. We now turn to a more formal discussion of the relationship between these cross-industry

and 47565. The TSUSA defines the specific tariff per gallon while the IDB records quantity by the barrel, where one barrel equals 42 gallons. Given the importance of oil imports, this issue affects more than 30% of annual imports in certain years. To facilitate simple calculation of specific tariff AVEs using unit values from the import data, our dataset records statutory specific tariffs in terms of the unit of quantity specified in the IDB data.

[18] There are 46 ports in our data.

[19] Examples include TSUSA product 64448 "Gold leaf, unmounted, over 11.40 square inches in area". The specific tariff in the TSUSA for this product is always specified in units of 1,140 square inches. In the IDB, on the other hand, the unit of quantity is 1,140 square inches between 1980 and 1983 and is square inches in all other years. As a result, the specific tariff calculated using the import data – i.e. duties divided by quantity – increases by a factor of 1,140 between 1979 and 1980 and then shrinks by a factor of 1,140 between 1983 and 1984. In such cases, our dataset overwrites the original quantity to ensure consistent measurement over time.

Table 1: Descriptive Statistics by Sector

Sector	Import Share	Line Share	Share Duty-Free	Share Specific	AVE_{1979}	ΔAVE 1972-1979	ΔAVE 1979-1988
Agriculture	0.120	0.182	0.208	0.606	0.073	-0.032	-0.015
Wood & Paper	0.064	0.059	0.260	0.385	0.050	0.004	-0.024
Textiles	0.047	0.155	0.045	0.315	0.159	-0.029	-0.064
Chemicals	0.152	0.158	0.214	0.521	0.055	-0.010	-0.013
Non-Metal Minerals & Prod.	0.035	0.070	0.178	0.333	0.098	-0.006	-0.033
Metals & Metal Products	0.486	0.206	0.145	0.254	0.069	0.004	-0.031
Misc. Manuf. Prod.	0.096	0.169	0.054	0.201	0.127	-0.008	-0.061

Notes: Data come from our combined import and tariff database. The first five columns use 1979 data only. "Import Share" and "Line share" are shares of total imports and all tariff lines accounted for by the sector. "Share Duty-Free" is the share of tariff lines in each sector not subject to tariffs. "Share specific" is the share of tariff lines in a sector subject to specific tariffs. AVE_{1979} is the good-level ad valorem equivalent tariff, defined as total duties divided by imports.

differences and the evolution of the aggregate AVE.

2.2 Aggregate AVE Tariff Changes

In exploring the behavior of the aggregate AVE throughout our sample, we emphasize two features of the data noted in Figure 1 above. First, we investigate the sharp decline in the aggregate AVE in the years prior to the Tokyo Round despite the absence of wholesale policy changes. Second, we provide an explanation for the relative stability of the AVE in the years following the agreement.

To begin, we express the aggregate AVE at time t as the sum of duties divided by the sum of import values, each aggregated across exporters v and five-digit TSUSA goods g:

$$AVE_t = \frac{\sum_g \sum_v D_{vgt}}{\sum_g \sum_v M_{vgt}} = \sum_g \sum_v s_{vgt} AVE_{vgt}. \tag{1}$$

Referring to an exporter-good pair as a "variety", equation (1) expresses AVE_t as the weighted average of variety-level $AVE_{vgt} = \tau_{vgt} + \frac{f_{vgt}}{p^*_{vgt}}$, with weights equal to the variety-level import shares $s_{vgt} = \frac{M_{vgt}}{\sum_g \sum_v M_{vgt}}$.

We decompose the change in AVE_t into its constituent parts by defining three mutually exclusive sets of imported varieties: varieties V^C continuing between time t_0 and t_1, exiting varieties V^X, and entering varieties V^N.[20] Specifically,

$$\begin{aligned}\Delta AVE_t \equiv & AVE_t - AVE_{t_0} \\ = & \underbrace{\sum_{v \in V^C} s_{vgt_0} \Delta AVE_{vgt}}_{\text{AVE Effect}} + \underbrace{\sum_{v \in V^C} AVE_{vgt_0} \Delta s_{vgt}}_{\text{Composition Effect}} + \underbrace{\sum_{v \in V^C} \Delta s_{vgt} \Delta AVE_{vgt}}_{\text{Covariance Effect}} \\ & + \underbrace{\sum_{v \in V^N} s_{vgt} AVE_{vgt} - \sum_{v \in V^X} s_{vgt_0} AVE_{vgt_0}}_{\text{Net Entry Effect}}. \end{aligned} \tag{2}$$

Equation (2) shows that changes in AVE_t can be separated into the sum of changes in variety-level AVE on continuing varieties conditional on the initial composition of imports (AVE effect), changes in the composition of imports on continuing varieties conditional on initial AVE (composition effect), the co-movement of variety-level AVEs and import shares (covariance effect), and changes due to the net entry of varieties (net entry effect). To shed light on the dynamics of the aggregate AVE in our sample, we use our database to construct this decomposition between 1972 and 1988, defining the base year as $t_0 = 1979$.

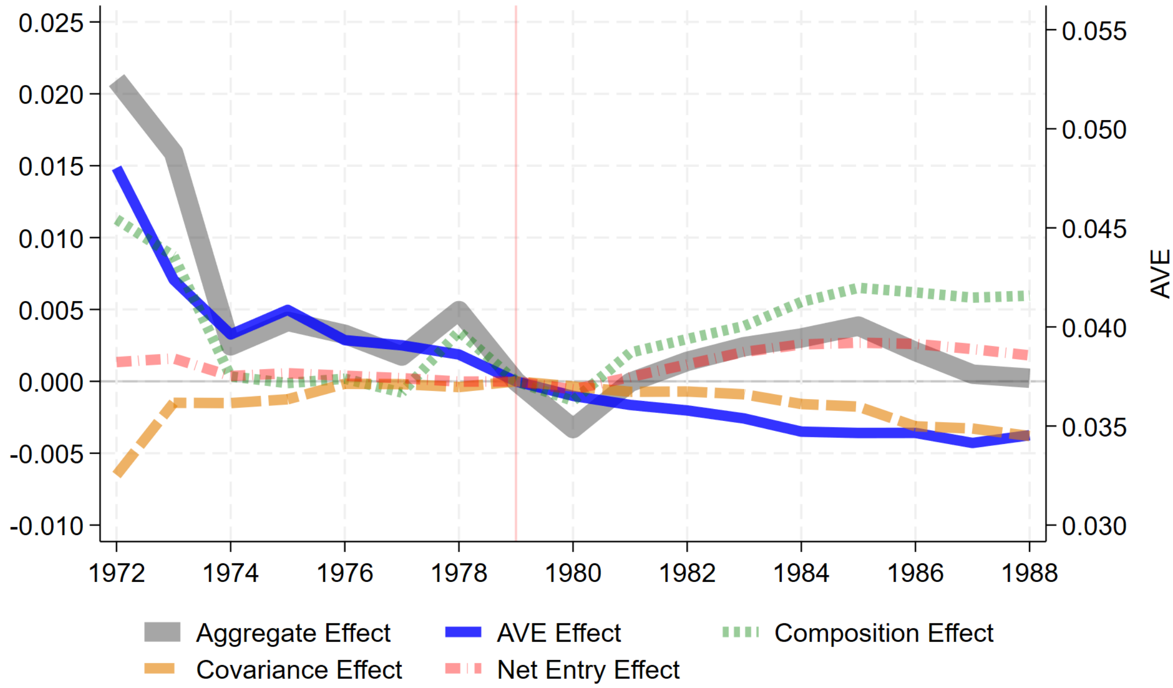
Figure 2 plots the decomposition.[21] The solid gray line depicts the aggregate AVE change ΔAVE_t, which closely matches that displayed in Figure 1. The AVE effect, depicted by the solid blue line, exhibits a substantial decline between 1972 and 1974, followed by a modest increase in 1975 and a continuous fall throughout the rest of our sample. The decline in aggregate AVE attributable to this channel between 1972 and 1979 amounts to an approximately 1.5 percentage point liberalization, despite this being a period in which column 1 Most Favored Nation (MFN) tariffs are unchanged by GATT rounds. The equivalent liberalization between 1979 and 1988, as Tokyo Round tariffs are phased in, is approximately one half of a percentage point. Overall, Figure 2 shows that variety-level AVE falls nearly continuously between 1972 and 1988.

Changes in the composition of imports also affect aggregate AVE dynamics. The dashed green line shows that the composition of imports during the 1970s shifts toward varieties

[20] Our decomposition analyses in equations (2) and (3) are similar to the decomposition in Crucini (1994), which explores the contributions of statutory tariff changes, rising import price indices, and changes in the relative price of imports to changes in AVEs between 1900 and 1940. However, because aggregation from the industry to the national level is not his focus, Crucini (1994) does not examine the composition or net entry effects. As mentioned below, variation in the composition effect explains over one-third of the variation in the aggregate AVE between 1972 and 1988.

[21] Appendix D details how the decomposition relates to sectoral trends. In particular, it explores the effects of the 1973-1974 and 1979-1980 oil crises.

Figure 2: Decomposition of Changes in Aggregate US AVE Tariff



Notes: Figure shows the decomposition of changes in the aggregate AVE, detailed in equation (2), with $t_0 = 1979$. The thick grey line plots ΔAVE_t. Each other line represents a component defined in equation (2). The left axis indicates the AVE change relative to 1979, while the right axis depicts the AVE. See main text for details.

with low 1979 AVEs, reducing the aggregate AVE by an additional percentage point. On the other hand, the shift of imports toward high 1979 AVE varieties in the 1980s masks the AVE effect related to the implementation of Tokyo Round tariff reductions. This pattern is consistent with the fact that, under the Swiss Formula, varieties experiencing greater post-1979 liberalization are those with initially higher tariffs. The covariance effect, depicted by the orange dashed line, similarly suggests a post-1979 shift towards varieties with falling AVE. Ultimately, changes in the composition of US imports mask the variety-level reductions in AVE in the wake of the Tokyo Round and, in turn, leave the aggregate AVE relatively constant between 1979 and 1988.[22]

Our discussion of Figure 2 underscores the fact that changes in the aggregate AVE do not necessarily reflect changes in tariff-line level tariffs. The masking of the AVE effect by

[22]The net entry effect, depicted by the dashed pink line, has little impact on explaining aggregate AVE changes.

the composition effect during the 1980s implementation of Tokyo Round tariff cuts provides a particularly stark example. This leaves unaddressed, however, whether variety-level AVE changes that underlie the AVE effect in equation (2) and Figure 2 are themselves an appropriate measure of liberalization. In other words, Figure 2 does not explain *why* variety-level AVEs change the way they do in this era. This issue is particularly salient for the years between 1972 and 1979, when the US tariff code is largely unchanged and variety-level AVEs fall dramatically. We explore this issue now.

2.3 Variety-Level Tariff Changes

Changes in variety-level AVEs can stem from any combination of changes in ad valorem tariffs τ_{vgt}, specific tariffs f_{vgt}, or tariff-exclusive prices p^*_{vgt}. As we emphasize below, these channels have quite different implications for import prices and volumes. It is thus important to understand the role of each in driving AVE changes throughout our sample.

To this end, we express changes in variety-level AVE between a base year t_0 and year t as

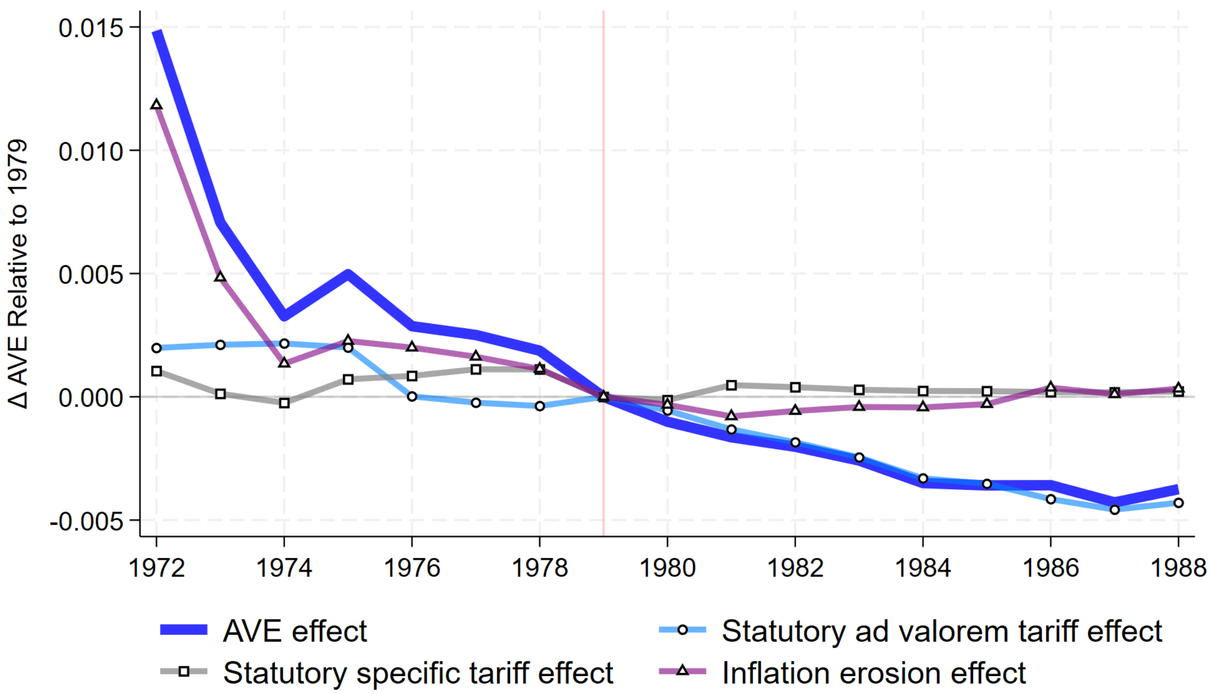
$$\begin{aligned}\Delta AVE_{vgt} &= \left(\tau_{vgt} + \frac{f_{vgt}}{p^*_{vgt}}\right) - \left(\tau_{vgt_0} + \frac{f_{vgt_0}}{p^*_{vgt_0}}\right) \\ &= \underbrace{\Delta\tau_{vgt}}_{\substack{\text{Statutory Ad Val.}\\ \text{Tariff Effect}}} + \underbrace{\frac{\Delta f_{vgt}}{p^*_{vgt_0}}}_{\substack{\text{Statutory Specific}\\ \text{Tariff Effect}}} - \underbrace{\frac{\Delta p^*_{vgt}}{p^*_{vgt_0}} STS_{vgt} AVE_{vgt}}_{\text{Inflation Erosion Effect}}, \end{aligned} \quad (3)$$

where $STS_{vgt} = \frac{f_{vgt}}{p^*_{vgt}\tau_{vgt}+f_{vgt}} \in [0,1]$ is the "specific tariff share," representing the proportion of variety-level tariff revenue coming from specific tariffs.[23] The statutory ad valorem tariff effect represents variety-level changes in statutory ad valorem tariffs. The statutory specific tariff effect represents the effect on variety-level AVE, evaluated at base period tariff-exclusive prices, of changes in the statutory specific tariff. Evaluating the specific tariff effect using base period prices purges it of any inflationary erosion component. Finally, the inflation erosion effect depends on the proportional change in the tariff-exclusive price $\frac{\Delta p^*_{vgt}}{p^*_{vgt_1}}$ and is higher when, in period t, specific tariffs are a more important component of the AVE or AVE is higher.

Figure 3 documents the relative importance of these channels in our sample. It does so

[23]The full derivation of this decomposition can be found in Appendix B.

Figure 3: Decomposition of Changes in Variety-Level AVE



Notes: Figure decomposes the AVE effect in equation (2) and Figure 2. It does so by decomposing changes in variety-level AVEs using equation (3) and aggregating using the variety-specific shares in equation (2). This aggregation ensures the share-weighted components from equation (3) sum to the AVE effect in equation (2) and Figure 2. See main text for details.

by decomposing changes in variety-level AVEs using equation (3) and aggregating using the variety-specific shares in equation (2). This implies that the aggregation yields the AVE effect in equation (2) and Figure 2.[24] The inflation erosion effect, depicted by the green line, accounts for 55%-85% of the AVE effect in years prior 1979, while statutory tariff changes play a minimal role.[25] Conversely, the post-1979 AVE effect is driven almost exclusively by changes in statutory ad valorem tariffs. Figure 3 thus underscores the fact that changes in AVEs can stem from dramatically different sources, even at the highly disaggregate exporter-

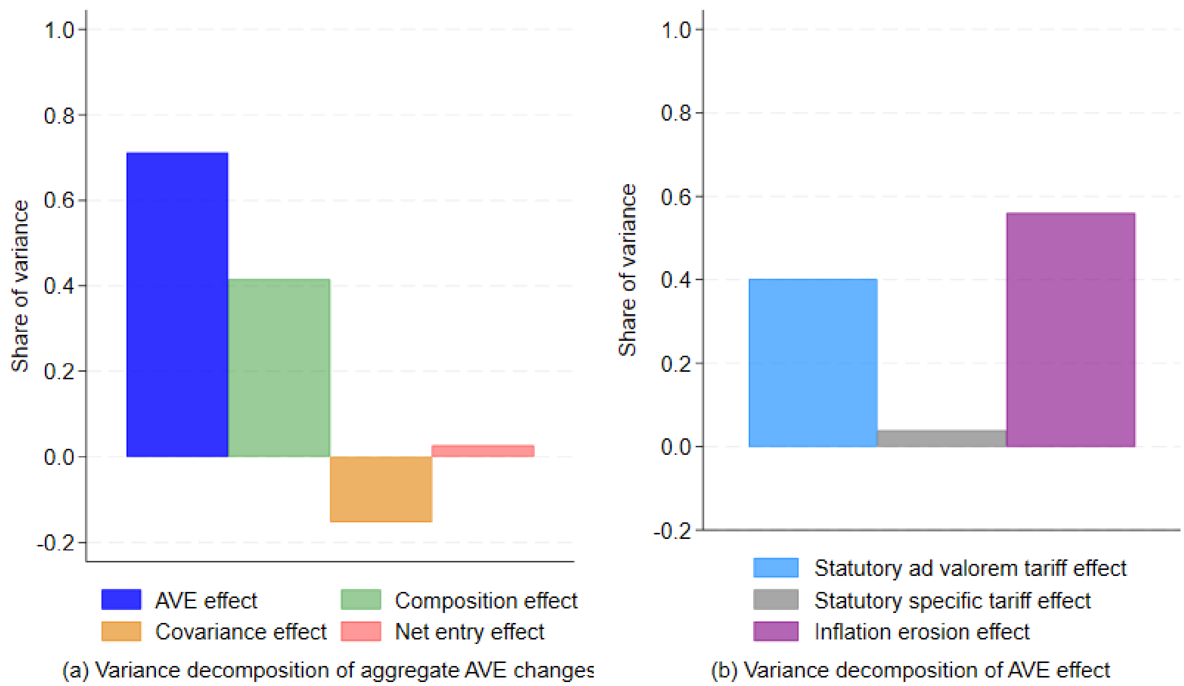
[24]There are numerous ways of decomposing the AVE effect. Instead of starting with AVEs at the exporter-good – i.e. variety – level, one could start with AVEs at rate type-by-good level with rate types corresponding to, e.g., column 1, column 2, and the various preferential tariff programs. This would highlight the relative importance of various rate types. We present such a decomposition in Appendix Figure A.2.

[25]The implementation of GSP in 1976 explains the bulk of the reduction in statutory ad valorem tariffs between 1975 and 1976. The temporary tariff suspensions on fuel in 1973-1974 and 1979-1980 account for the fluctuations in the statutory specific tariff effects in the 1972-1981 period. See Appendix D for further discussion of the sectors driving the various components of Figure 2.

tariff-line level and within relatively narrow periods of time.

In sum, nearly continuous reductions in variety-level AVE characterize the years surrounding the successful completion of Tokyo Round negotiations. However, changes in the composition of imports mask this decline in the aggregate post-Tokyo AVE. Further, the nature of the variety-level decline changes dramatically over time. Prior to 1979, inflation combines with pervasive specific tariffs to erode variety-level AVEs, while statutory ad valorem tariff cuts dominate the post-1979 years.

Figure 4: Sources of Variation in Aggregate AVE Changes



Notes: Figure decomposes the variance of the Aggregate Effect in Figure 2 and the AVE effect in Figure 3. It decomposes these into the variance of the constituent components in the respective decompositions in Figure 2 and Figure 3, and the pairwise covariance between each of the constituent components. For illustration, the covariance term between two constituent effects is equally allocated between the two effects.

We summarize the importance of the respective sources of variation in AVEs in Figure 4. Panel (a) reports the share of the variation in aggregate AVE changes accounted for by each of the components described above.[26] The panel indicates that variation in the AVE

[26]For any random variable $X = \sum_{i=1}^{N} X_i$, the variance of X can be written as $\text{var}(X) = \sum_{i=1}^{N} \text{var}(X_i) + \sum_{i=1}^{N} \sum_{j=1, j \neq i}^{N} \text{cov}(X_i, X_j)$. Note that each covariance term $\text{cov}(X_i, X_j)$ appears as $2\text{cov}(X_i, X_j)$ when expanding the summation of covariances. For illustration, we allocate one of these covariances to the "X_i"

effect and the composition effect explain 71% and 42% of the annual variation in aggregate AVE changes, respectively. Panel (b) depicts the sources of annual variation in the AVE effect. The inflation erosion effect and the statutory ad valorem tariff effect explain 56% and 40% of variation in the AVE effect. That is, consistent with Figures 2 and 3, while the AVE effect accounts for the greatest share of variation in aggregate AVE changes, compositional effects also play an important role. Further, approximately 60% of the annual variation in the AVE effect is driven not by variation in statutory tariff changes, but by variation in the inflation erosion effect. These results underscore the fact that the aggregate AVE is not a policy variable: less than 30% of the variation in annual aggregate AVE changes is driven by variation in statutory tariff changes.[27] We now explore the importance of this distinction, showing that statutory tariff changes and inflation erosion have dramatically different implications for imports and prices.

3 A Quantitative Framework with Specific Tariffs

Understanding the evolution of US tariffs in the 1970s and 1980s requires one to distinguish between various sources of AVE changes. Here, we quantify the importance of these sources for domestic price indices and the interpretation of counterfactual results. Section 3.1 embeds specific tariffs in a standard CES framework and Section 3.2 extends existing hat algebra techniques to this setting.[28]

3.1 Model

On the demand side, a representative consumer has a three-tier utility function with time-varying expenditure E_t allocated at the upper tier across a domestic composite good q_{Dt} and an importable composite good q_{Mt}. To focus our analysis on the distinction between specific and ad valorem tariffs, we do not model the domestic production side of the economy.

effect and one to the "X_j" effect.

[27] That is, 71% of the variation comes from the AVE effect, and 40% of this, or less than 30% of the total variation, comes from statutory tariffs. In Appendix G, we conduct similar exercises to those documented here for the years 1989-2016. While the role of specific tariffs in explaining annual variation in AVEs has declined over time, we show that compositional effects remain important in modern data.

[28] Appendices E.1 and E.2 provide full descriptions of, respectively, the hat algebra and exact hat algebra representations of the model.

Instead, we assume a Cobb-Douglas upper tier of utility with expenditure shares α and $1-\alpha$:

$$u\left(q_{Dt}, q_{Mt}\right) = q_{Dt}^{\alpha} q_{Mt}^{1-\alpha}. \tag{4}$$

The middle tier defines consumption of the importable composite good as a CES aggregator across industries I with an elasticity of substitution $\gamma > 1$:

$$q_{Mt} = \left[\sum_{i\in I} q_{it}^{(\gamma-1)/\gamma}\right]^{\gamma/(\gamma-1)}. \tag{5}$$

Finally, the lower tier defines consumption of composite importable industry q_{it} as a CES aggregate across exporters v and goods g within industry i with an elasticity of substitution $\sigma > 1$ and a time-varying, variety-specific taste shock b_{vgt}:[29]

$$q_{it} = \left[\sum_{\{v,g\}\in i} b_{vgt}^{1/\sigma} q_{vgt}^{(\sigma-1)/\sigma}\right]^{\sigma/(\sigma-1)}. \tag{6}$$

Tariffs create a wedge between the tariff-exclusive price received by the foreign exporter, p^*_{vgt}, and the tariff-inclusive price paid by domestic consumers, p_{vgt}:

$$p_{vgt} = p^*_{vgt}\left(1 + AVE_{vgt}\right), \tag{7}$$

where AVE_{vgt} is the variety-level AVE tariff:

$$AVE_{vgt} \equiv \tau_{gt} + \frac{f_{gt}}{p^*_{vgt}}. \tag{8}$$

Export supply is given by

$$p^*_{vgt} = \exp\left(\eta_{vgt}\right) q_{vgt}^{\omega}, \tag{9}$$

where η_{vgt} is a time-varying, variety-specific technology shock and $\omega \geq 0$ is the inverse export supply elasticity.

Equilibrium variety-level prices equate import demand and export supply. After a log transformation, the variety-level import demand and inverse export supply curves can be

[29] We define i at the five-digit TSUSA level in the counterfactuals and either four-digit or five-digit TSUSA goods in the empirics. Throughout, we define g at the five-digit TSUSA level.

written as

$$\ln q_{vgt} = -\sigma \ln p_{vgt} + \phi_{it} + u_{D,vgt} \tag{10}$$

$$\ln p^*_{vgt} = -\omega \ln q_{vgt} + u_{S,vgt}, \tag{11}$$

where $\phi_{it} = \ln E_{it} + (\sigma - 1) \ln P_{it}$ combines the industry-level expenditure and price index and $u_{D,vgt} = \ln b_{vgt}$ and $u_{S,vgt} = \eta_{vgt}$ represent the structural import demand and export supply shocks, respectively.[30] Equilibrium prices and quantity are then

$$\ln p^*_{vgt} = -\frac{\sigma\omega}{1+\sigma\omega} \ln(1 + AVE_{vgt}) + \varphi_{p^*,it} + u_{p^*,vgt} \tag{12}$$

$$\ln p_{vgt} = \frac{1}{1+\sigma\omega} \ln(1 + AVE_{vgt}) + \varphi_{p,it} + u_{p,vgt} \tag{13}$$

$$\ln q_{vgt} = -\frac{\sigma}{1+\sigma\omega} \ln(1 + AVE_{vgt}) + \varphi_{q,it} + u_{q,vgt}, \tag{14}$$

where the $\varphi_{\cdot,it}$ terms combine the industry-level expenditure and price index while the $u_{\cdot,vgt}$ terms combine the structural import demand and export supply shocks.

Our counterfactual estimation requires elasticity estimates. In Appendix E.3, we combine equations (10)-(14) and the approach developed by Zoutman et al. (2018) and used in a trade context by Fajgelbaum et al. (2020) with an instrument for tariff changes based on the Swiss Formula (Greenland et al., 2026) to derive import demand and export supply elasticities. In our preferred specification we obtain an import demand elasticity of -3.5, well within the range of estimates found in the literature, and an export supply elasticity that is not statistically different from zero. This implies that tariff reductions are passed through entirely into lower tariff-inclusive prices, with no accompanying effect on tariff-exclusive prices. This small open economy interpretation is analogous to the results of Fajgelbaum et al. (2020), who find that US tariff increases during the first US-China trade war had no impact on tariff-exclusive prices. The result also significantly simplifies the counterfactual analysis in Section 4.

3.2 Hat Algebra

The counterfactual analysis in Section 4 to follow employs the exact hat algebra detailed in Appendix E.2, which defines $\hat{x} \equiv \frac{x'}{x}$ as the counterfactual value x' relative to the initial value

[30] The industry-level price index is $P_{it} = \left(\sum_{\{v,g\}\in i} b_{vgt} p_{vgt}^{1-\sigma}\right)^{1/(1-\sigma)}$.

x for a variable of interest. To capture the key intuition underlying this approach, we focus here on the first order approximation of the model, commonly referred to as hat algebra, defining $\hat{x} \equiv d\ln x = \frac{dx}{x}$.

To formalize our discussion surrounding the effects of changes in AVEs, we write the proportional change in $1 + AVE_{vgt}$ as

$$\widehat{(1 + AVE}_{vgt}) = \frac{dAVE_{vgt}}{1 + AVE_{vgt}} \tag{15}$$

$$= \frac{d\tau_{gt} + df_{gt}\frac{1}{p^*_{vgt}}}{1 + AVE_{vgt}} - STS_{vgt}\frac{AVE_{vgt}}{1 + AVE_{vgt}}\hat{p}^*_{vgt}. \tag{16}$$

This reduces to $\frac{d\tau_{gt}}{1+\tau_{gt}}$ in the absence of specific tariffs, as in, e.g., Fajgelbaum et al. (2020). However, specific tariffs modify this standard expression in two ways. First, they introduce a second statutory tariff component, $df_{gt}\frac{1}{p^*_{vgt}}$, in addition to the ad valorem statutory component, $d\tau_{gt}$. Second, they generate the term capturing the response of AVE_{vgt} to the tariff-exclusive price change $\hat{p}^*_{vgt}$ via the (incomplete) pass-through coefficient $STS_{vgt}\frac{AVE_{vgt}}{1+AVE_{vgt}} \in [0, 1)$.[31]

With an eye toward aggregate price indices, we write the proportional change in the tariff-inclusive price as the sum of the proportional changes in the AVE and the tariff-exclusive price. Using equation (16), this can be written as

$$\hat{p}_{vgt} = \widehat{(1 + AVE}_{vgt}) + \hat{p}^*_{vgt} \tag{17}$$

$$= \frac{d\tau_{gt} + df_{gt}\frac{1}{p^*_{vgt}}}{1 + AVE_{vgt}} + \left(1 - STS_{vgt}\frac{AVE_{vgt}}{1 + AVE_{vgt}}\right)\hat{p}^*_{vgt}. \tag{18}$$

An important point that emerges from equations (16)-(18) is that, in the presence of specific tariffs, one cannot infer whether tariff-inclusive prices rise or fall following a reduction in AVE. Absent specific tariffs, equation (16) implies that AVE movements depend only on statutory tariff changes. In turn, equation (17) implies that tariff-inclusive prices fall as AVE falls. This remains true when *statutory* specific tariffs fall. However, in the presence of specific tariffs the AVE also depends on tariff-exclusive prices. Given the incomplete pass-through from tariff-exclusive prices to AVE in equation (16), equation (18) implies that tariff-inclusive prices *increase* in the face of reductions in AVE_{vgt} driven by increases

[31] An implication of this incomplete pass-through coefficient is that specific tariff policy affects the volatility of tariff-inclusive prices. We explore this issue in Greenland et al. (2025).

in p^*_{vgt}. This is due to the fact that inflationary erosion of the AVE in response to rising tariff-exclusive prices only partially offsets the direct effect that such price increases exert on tariff-inclusive prices.

The importable price index, then, is an import-share-weighted average of changes in good-level price indices:[32]

$$\hat{P}_{Mt} = \sum_{g=1}^{G} s_{gt} \hat{P}_{gt}. \tag{19}$$

Good-level price indices, in turn, are import-share-weighted averages of the change in variety-level tariff-inclusive prices and elasticity-adjusted demand shocks:[33]

$$\hat{P}_{gt} = \sum_{v=1}^{V_g} s_{vgt} \left(\hat{p}_{vgt} - \frac{1}{\sigma - 1} \hat{b}_{vgt} \right). \tag{20}$$

As one cannot infer whether tariff-inclusive prices rise or fall when AVE falls in an environment with specific tariffs, equations (19) and (20) imply that one cannot infer whether the aggregate price index rises or falls following reductions in AVE.

Specific tariffs also complicate the relationship between AVEs and imports. To see this, we write proportional changes in import value as

$$\begin{aligned} \hat{m}_{vgt} &= \hat{p}^*_{vgt} + \hat{q}_{vgt} \\ &= \frac{1 + \omega \rho_{vgt}}{1 + \sigma \omega} \left[-\sigma \frac{d\tau_{gt} + df_{gt} \frac{1}{p^*_{vgt}}}{1 + AVE_{vgt}} + \hat{b}_{vgt} + \hat{E}_{gt} + (\sigma - 1) \hat{P}_{gt} \right] - \frac{\sigma - \rho_{vgt}}{1 + \sigma \omega} \tilde{\eta}_{vgt}, \end{aligned} \tag{21}$$

which depends on three underlying shocks: statutory tariff changes $d\tau_{vgt}$ and df_{vgt}, import demand shocks $\hat{b}_{vgt}$, and foreign export supply shocks $\tilde{\eta}_{vgt}$. Specific tariffs enter equation (21) through statutory tariff changes and ρ_{vgt}.[34] As emphasized above, in the absence of specific tariffs, changes in AVE stem only from statutory tariff cuts. In this case, equation (21) implies that the value of imports increases as AVE falls. In the presence of specific tariffs, however, the mapping from AVE to import values implied by equation (21) is ambiguous. On

[32] In the context of equation (5), we hereafter define the lower nest at the five-digit TSUSA good level.

[33] The change in tariff-inclusive prices, $\hat{p}_{vgt}$, also depends on demand shocks. These shocks enter the good-level price index separately to their influence on $\hat{p}_{vgt}$ because positive (negative) shocks reduce (increase) the expenditure required to "purchase" a unit of utility.

[34] Note that $\rho_{vgt} = \left[1 + \frac{\sigma}{1+\sigma\omega} STS_{vgt} \frac{AVE_{vgt}}{1+AVE_{vgt}}\right] \left[1 + \frac{\sigma\omega}{1+\sigma\omega} STS_{vgt} \frac{AVE_{vgt}}{1+AVE_{vgt}}\right]^{-1} \geq 1$ and reduces to $\rho_{vgt} = 1$ in the absence of specific tariffs.

one hand, the value of imports increases with AVE reductions driven by statutory tariff cuts or by increases in tariff-exclusive prices stemming from positive demand shocks ($\hat{b}_{vgt} > 0$). On the other hand, if $\sigma > \rho_{vgt}$, the value of imports decreases with AVE reductions driven by higher tariff-exclusive prices stemming from negative supply shocks ($\tilde{\eta}_{vgt} > 0$). Thus, as with tariff-inclusive prices, one cannot infer whether the value of imports rises or falls from changes in AVE alone.

The relationships between tariffs, prices, and imports are fundamental in international trade. Our analysis emphasizes that specific tariffs imply that these relationships depend on the source of variation in AVE. Analyses that conflate statutory tariff changes with AVE movements driven by underlying supply shocks, for example, risk misstating the effects of the tariffs on prices and imports not only in magnitude but in direction. We now demonstrate the importance of this general point in the context of the Tokyo Round.

4 Statutory and Inflationary Tariff Counterfactuals

In this section, we conduct a series of counterfactual exercises aimed at quantifying the effects of US tariff policy during this period. We first decompose the effect of changes in variety-level AVE on US import prices into their statutory and inflationary components. We then demonstrate empirically that changes in AVE are not sufficient for determining changes in aggregate price indices and explore the implications of this result in the context of the Tokyo Round. Finally, we explore sectoral heterogeneity along these dimensions.

Counterfactual analysis of AVEs in the presence of specific tariffs requires assumptions regarding the relationship between unobserved demand and supply shocks, as AVEs depend on prices, which in turn depend on these shocks. For example, counterfactual imports depend on whether price-induced AVE changes reflect exporter productivity shocks or factors that simultaneously affect export supply and import demand, such as quality upgrading or forces driving global inflation. Given the fact that we do not model the possible correlation between import demand and export supply shocks, the analysis to follow assumes that these shocks are uncorrelated. However, alternative settings may of course require different assumptions. This further underscores the point that analysis of AVE requires researchers to understand the source of AVE variation in their setting.

4.1 Statutory vs. Inflationary Liberalization

As detailed above, AVEs fall in the years surrounding the Tokyo Round for both statutory and inflationary reasons. To assess the contribution of these separate channels to changes in import price indices, we consider counterfactuals that separate disaggregate AVE changes into their statutory and inflationary components. Our first counterfactual considers a scenario in which all tariffs are ad valorem and are left unchanged after 1972.[35] The difference between this counterfactual and the observed data corresponds to an estimate of the changes in the import price index in the absence of all liberalization, statutory or inflationary, while allowing tariff-exclusive prices to evolve as observed. Our second counterfactual allows for statutory liberalizations only.[36] The difference between this exercise and the observed data thus implies the magnitude of the liberalization due to inflationary channels.[37]

The combined height of the blue and green bars in Figure 5 shows the cumulative effect on the importable price index under our first counterfactual. This represents the estimated effect on the importable price index of removing all observed changes in AVEs. The figure reveals that, by the end of our sample, the importable price index would have been approximately 2 percentage points higher in the absence of such changes. Aside from a temporary fluctuation in the early-mid 1970s driven by the statutory changes in fuel-related tariffs surrounding the 1973 oil crises, the aggregate effect grows continuously throughout the sample.

Our second counterfactual differs from the first by allowing for AVE changes driven exclusively by observed statutory tariff changes. Comparing this to the observed data thus reveals the effect of AVE inflationary erosion on the importable price index. The blue bars in Figure 5 depict these inflationary effects. In turn, the residual green bars represent the effect of changes in statutory tariffs.[38]

The figure underscores the importance of inflation-driven liberalization. The height of

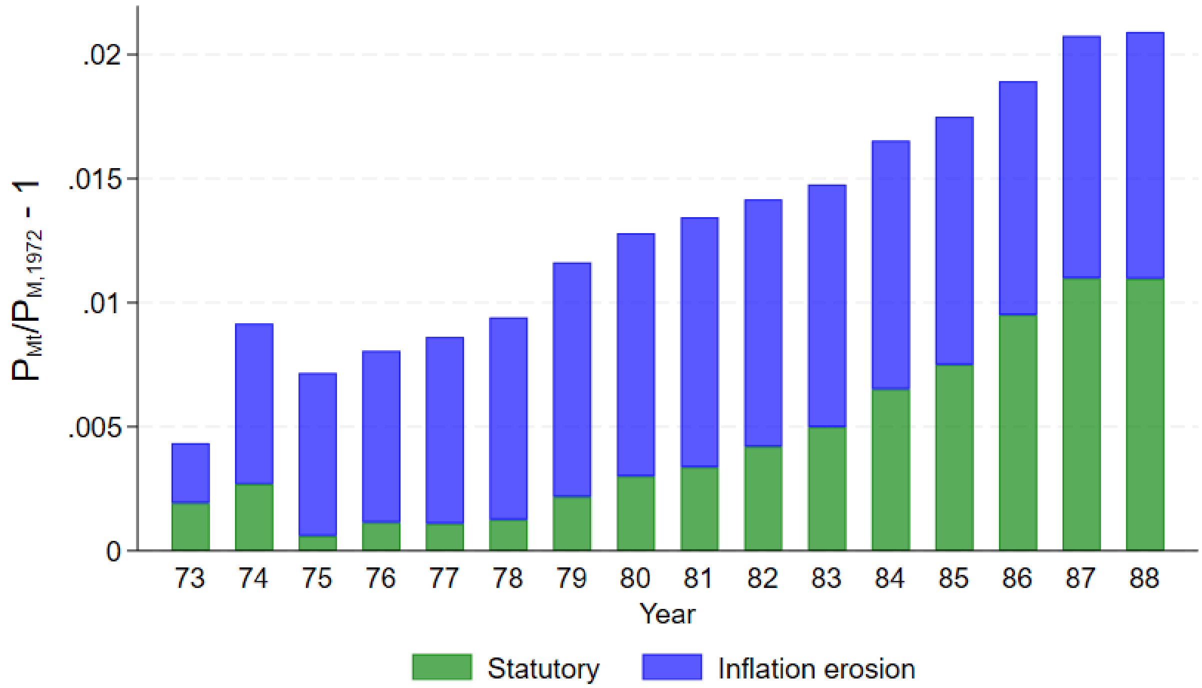
[35]To do so, we impose counterfactual 1972 ad valorem tariffs equal to their observed 1972 AVE values.

[36]This counterfactual continues to impose counterfactual 1972 ad valorem tariffs equal to their observed 1972 AVE values. However, unlike the first counterfactual, ad valorem tariffs in this case evolve with observed changes in statutory tariffs. For changes in statutory specific tariffs, the counterfactual change in the ad valorem tariff is the AVE of the specific tariff change using 1979 tariff-exclusive prices, implying that it solely reflects statutory tariff changes.

[37]To implement these counterfactuals, we impose the counterfactual tariff changes on the observed data in each year. This effectively creates a counterfactual time series in which all tariffs were ad valorem and either did not change after 1972 – i.e., the first counterfactual – or only changed due to statutory tariff changes.

[38]The results are nearly identical if the second counterfactual instead removes statutory tariff changes and hence leaves the remainder as indirectly revealing the impact of inflation erosion. The two approaches would be equivalent if the model was linear, but the non-linearity in the model implies that they are not identical.

Figure 5: Cumulative Importable Price Index Increase Removing Sources of AVE Decline



Notes: Figure presents exact hat counterfactual estimates of the cumulative effect of changes in AVEs on the importable price index using equation (E.27) in Appendix E. Green (blue) bars represent the percentage increase in the importable price index if statutory tariff changes (inflation erosion of specific tariffs) had not happened. Combined height represents effect in the absence of any AVE changes.

the blue bar in 1974, for example, reveals that the importable price index would have been roughly 0.6 percentage points higher had rising tariff-exclusive prices not eroded the protection afforded by specific tariffs. This magnitude increases throughout the 1970s and peaks at 0.94 percentage points following the second oil crisis in 1981. Apart from temporary tariff suspensions on fuel imports and the implementation of GSP, nearly all changes in the importable price index during the 1970s are driven by the inflationary erosion of specific tariffs.

In contrast, the inflationary component remains largely static throughout the 1980s, while the role of statutory tariffs increases.[39] Consistent with the linear phase-in of Tokyo Round tariffs between 1980 and 1987, the statutory component grows steadily throughout

[39]As we discuss in Appendix F, the US dollar appreciated sharply during the early-mid 1980s. Absent this appreciation, inflationary erosion of specific tariffs would have continued due to rising foreign currency tariff-excusive prices. By 1988, this would have contributed an additional 1 percentage points to the decline in the importable price index.

the decade and reaches a peak of 1.1 percentage points in 1987. By the end of our sample, the combined liberalization generated by the inflationary erosion of specific tariffs and statutory tariff cuts reduces the importable price index by approximately 2 percentage points, with the two channels contributing essentially equal amounts.

4.2 Interpreting Changes in AVE

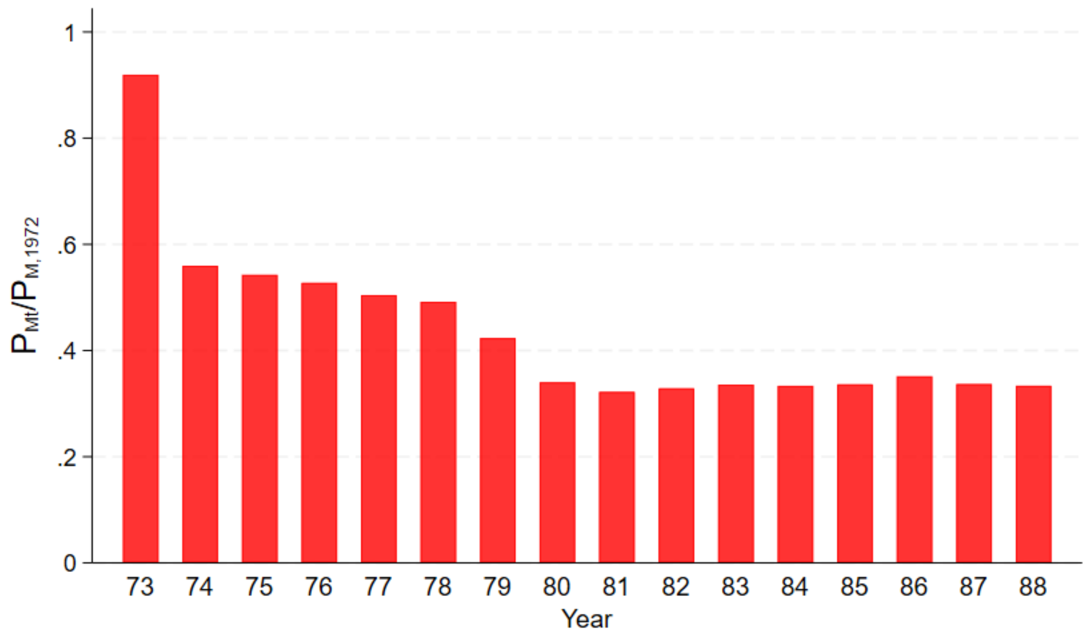
The counterfactual exercises above describe the relative contribution of statutory and inflationary channels to tariff liberalization, and the associated impact of these channels on the importable price index through their effect on AVE. They do not, however, capture the overall effect on aggregate price indices of the underlying shocks driving the changes in AVE. As discussed in Section 3.2, such shocks may affect the price index directly, potentially in a way that offsets the effect of changes in AVE. In particular, if negative supply shocks drive the inflationary channel depicted in Figure 5, then rising price levels may accompany falling AVEs. More broadly, AVE changes conflate changes in equilibrium objects – i.e., tariff-exclusive prices – with exogenous policy changes. One can thus not interpret the effects of AVE changes on outcomes of interest while ignoring the source of those changes.

We address this issue by conducting a counterfactual in which we remove observed changes in tariff-exclusive prices and thus capture the direct effect of variety-level price movements on the aggregate price index. To isolate this direct effect, we continue to assume that all tariffs are ad valorem, which implies that price changes do not affects AVEs. Given the fact that our empirical estimates from Appendix E.3 suggest a perfectly elastic export supply curve, our model interprets the observed changes in tariff-exclusive prices as export supply shocks – i.e., $\hat{p}^*_{vgt} = \hat{\eta}_{vgt}$. We thus conduct the counterfactual by removing implied export supply shocks from the data. Figure 6 presents the results of this exercise, again reporting the cumulative impact on the importable price index over time. Comparing this figure to the blue bars in Figure 5, which depict the effect of inflation on the importable price index through its effect on AVEs, allows us to assess the overall implications of price movements – both directly and through inflationary erosion – in this era.

Like Figure 5, Figure 6 shows the importance of inflation in the era: the importable price index would have been nearly 45% lower in 1974 than 1972 absent the model-implied 1973 export supply shocks and, especially, the oil-crisis-related 1974 export supply shocks.[40]

[40] See Appendix D for further discussion on the role of oil throughout our sample period.

Figure 6: Cumulative Importable Price Index Decline Removing Export Supply Shocks



Notes: Figure presents exact hat counterfactual estimates of the cumulative effect of foreign export supply shocks on the importable price index using equation (E.27) in Appendix E. Bars represent the magnitude of decline had the observed export supply shocks not happened. Our elasticity estimate of $\omega_g = 0$ implies the model interprets observed changes in tariff-exclusive prices as foreign export supply shocks.

Moreover, these effects continue to grow through the remainder of the decade. By the end of the second oil crisis in the early 1980s, the importable price index would have been approximately 70% lower absent the export supply shocks of the prior decade. Comparing these magnitudes to those in Figure 5 underscores the point that AVE cannot be used in isolation to assess the consequences of trade liberalization for aggregate price indices. Inflationary erosion of specific tariffs in the years surrounding the Tokyo Round does indeed reduce AVEs, and the reduction in AVEs does indeed reduce prices. But, the underlying shocks that drive inflation increase prices, and these direct effects are an order of magnitude larger than the indirect ones operating through AVE.

This result illustrates an important point: counterfactual effects of AVE changes depend heavily on the underlying counterfactual of interest. Figure 5 asks how the inflationary erosion of AVEs affects the importable price index *conditional* on the underlying tariff-exclusive price shocks, and indicates that aggregate prices fall as AVE falls. One could alternatively interpret this as reflecting the gains of having the observed specific tariffs rather than ad valorem tariffs. However, a counterfactual with no inflationary erosion of AVEs is implicitly one without underlying price shocks. To this end, Figure 6 asks instead how the importable price index would have changed absent such shocks. This figure implies that prices *rise* as AVEs fall – not because of the tariff liberalization itself, but because of the underlying shocks that drive it.

The above highlights the importance of distinguishing between sources of changes in AVEs for structural trade models exploring the relationship between AVEs, prices, and imports. A similar point applies in reduced-form settings as well, including estimating the impact of AVE changes on imports. To this end, Table 2 provides estimates of the effect of AVE changes on model-implied changes in the value of imports. Specifically, we estimate

$$\Delta \ln(m_{vgt}) = \beta \Delta \ln(1 + AVE_{vgt}) + \gamma_{gt} + \varepsilon_{vgt}, \tag{22}$$

where γ_{gt} are five-digit TSUSA good-by-year fixed effects. Standard errors are clustered by exporter and five-digit TSUSA good. As a benchmark, column 1 estimates the effect of observed statutory changes in AVEs on the model-implied growth in import value under the assumption that $\sigma = -3.5$ and $\omega = 0$. In doing so, we mechanically recover the imposed elasticity of $\sigma = 3.5$.

Table 2: Counterfactual Import Growth by Source of AVE Change

	(1)	(2)	(3)
$\Delta \ln(1 + AVE_{vgt})$	-3.50***	17.70***	-3.50***
	(0.00)	(1.71)	(0.00)
$\Delta \ln(p^*_{vgt})$			-2.50***
			(0.00)
Statutory tariff shocks	Yes	No	No
Price shocks	No	Yes	Yes
N	535,436	535,436	535,436
R^2	1.00	0.15	1.00

Notes: Table presents estimates of equation (22). Dependent variable is model-implied change in imports Δm_{vgt} with $\sigma = 0$ imposed. All columns include five-digit TSUSA by year fixed effects. Two-way standard errors clustered by exporter and five-digit TSUSA good.
$^{*}p < .1, ^{**}p < .05, ^{***}p < .01$

Columns 2 and 3 highlight the endogeneity problem associated with variation in AVEs driven by variation in prices. In these columns, we fix statutory tariffs, which implies that AVE changes are driven solely by price changes. As above, our model interprets these as export supply shocks given $\omega = 0$. While the column 2 coefficient estimate is highly statistically significant, it is the wrong sign. Intuitively, this positive relationship between changes in AVEs and imports reflects the fact that export supply shocks increase tariff-exclusive prices, leading to reduced AVEs and imports. However, when we control for growth in tariff-exclusive prices in column 3, we again recover the imposed elasticity of $\sigma = 3.5$. By controlling for prices, the specification accounts for the fact that export supply shocks lead to two offsetting effects on imports: a direct effect whereby higher tariff-exclusive prices reduce imports and an indirect effect whereby the higher tariff-exclusive prices lead to lower AVEs and, in turn, higher imports.[41]

These results have important implications for the measurement of exposure to trade liberalization. The literature has largely adopted AVE changes as its preferred measure, and it is easy to see why. AVEs are easy to calculate, can be aggregated to the desired level, and, in the absence of specific tariffs, lead to an intuitive relationship between prices, imports, and welfare. However, specific tariffs complicate these relationships because their presence implies the relationships – both in magnitude and sign – depend on the underlying shocks driving AVE changes. This underscores the importance of statutory tariff-line data. Such data allow one to distinguish between policy-driven changes in AVE tariffs versus those driven by inflation, which is necessary to correctly identify the effects of interest.

4.3 Sectoral Heterogeneity

Our results also have implications for the interpretation of the consequences of trade liberalization across sectors. Heterogeneity in the reliance on both tariffs in general and specific tariffs in particular generates heterogeneity in the magnitude and nature of tariff liberalization. We demonstrate this point in Table 3, in which we present counterfactual estimates akin to those in Figures 5 and 6 separately for each one-digit TSUSA sector as well as separately, and combined, for the pre-Tokyo and post-Tokyo periods. Columns (1), (4), and (7) report the counterfactual effects of removing observed statutory tariff cuts while columns (2), (5), and (8) report the counterfactual effects of removing inflation erosion of specific tariff AVEs. These correspond to Figure 5. Columns (3), (6), and (9) report the effects of reversing the observed price changes on specific tariff goods that drive inflation erosion and correspond to Figure 6.

[41]This result depends on a zero export supply elasticity. Otherwise, in the presence of specific tariffs, import demand shocks also affect the AVE through their effect on the equilibrium tariff-exclusive price. In such cases, as long as export supply is not perfectly inelastic, the error term will contain import demand shocks.

Table 3: Counterfactual Changes in Importable Price Index

	Pre-Tokyo (1972–1979)			Post-Tokyo (1979–1988)			Full Sample Period		
	(1) Stat. Tariffs	(2) Inflation Erosion	(3) Supply Shocks	(4) Stat. Tariffs	(5) Inflation Erosion	(6) Supply Shocks	(7) Stat. Tariffs	(8) Inflation Erosion	(9) Supply Shocks
Agriculture	0.001	0.019	0.438	0.006	0.008	0.719	0.007	0.027	0.315
Wood & paper	0.001	0.000	0.477	0.009	0.000	0.825	0.010	0.000	0.393
Textiles	0.001	0.030	0.352	0.044	0.002	0.634	0.044	0.032	0.223
Chemicals	0.011	0.014	0.125	0.000	−0.002	1.012	0.011	0.012	0.127
Non-metal minerals & prod.	0.002	0.002	0.451	0.024	0.000	0.590	0.026	0.003	0.266
Metals & metal products	0.000	0.003	0.330	0.009	0.000	0.576	0.010	0.003	0.190
Misc. manuf. prod.	0.005	0.002	0.428	0.016	−0.001	0.324	0.021	0.001	0.139

Notes: Using the counterfactuals depicted in Figures 5 and 6, the table reports cumulative changes in sector-level price indices over the period indicated. The sector-level importable price index is defined using equation (E.27) in Appendix E as $\hat{P}_{st} = \prod_{g \in s} \hat{P}_{gt}^{\tilde{w}_{gt}}$ for one-digit TSUSA sector s where the weights $\tilde{w}_{gt}$ normalize the weights w_{gt} such that the weights $\tilde{w}_{gt}$ sum to 1. Appropriate sector-level weights can then be used to aggregate the sector-level price indices $\hat{P}_{st}$ to the importable price index $\hat{P}_{Mt}$.

There is substantial heterogeneity across sectors and over time in the extent and type of liberalization. The inflationary erosion of specific tariffs drives the 1970s liberalization.[42] Absent the erosion of specific tariff AVEs, column (2) shows the importable price index would have been higher 1.9% higher in "Agriculture", 3% higher in "Textiles", and 1.4% higher in "Chemicals". However, the importable price index would have been at least 50% lower in each of these sectors in the absence of the supply shocks driving higher tariff-exclusive prices. In contrast, statutory tariff cuts drive the 1980s liberalization.[43] Absent these statutory tariff cuts, the importable price index would have been 4.4% higher in "Textiles", 2.6% higher in "Non-metal Minerals & Products", and 2.1% higher in "Miscellaneous Manufactured Products".

Importantly, this discussion highlights considerable variation across sectors in the nature of liberalization, with accompanying implications for prices and trade. The counterfactual increase in the importable price index absent any tariff liberalization is similar in "Agriculture" and "Non-metal Minerals and Products", at 3%-3.5%. However, the source of AVE liberalization differs dramatically across these sectors. Roughly 80% of the effect in "Agriculture" is due to inflationary erosion, while approximately 90% of the effect in "Non-metal Mineral & Products" is due to statutory tariff changes. On the one hand, the tariff liberalization in "Non-metal Minerals & Products" thus generates higher imports and lower prices,

[42] The exception is "Chemicals" where tariff suspensions on oil imports in 1979 represent temporary statutory tariff cuts. See Appendix D for more details.

[43] An exception is "Agriculture", in which the importable price index would have been 0.7% higher in the absence of inflation erosion of specific tariff AVEs.

as illustrated by column (1) of Table 2 and Figure 5. On the other hand, the inflation-driven liberalization in "Agriculture" generates lower imports, as illustrated by column (2) of Table 2 and Figure 6. In the presence of specific tariffs, all liberalizations are not created equal.

In Appendix G, we show that sectoral heterogeneity in the nature of AVEs remains important in modern settings. In particular, we show that specific tariffs account for 30 to 50% of the cross-sectional variation in good-level AVEs between 1989 and 2017, with the majority of this variation driven by differences in prices rather than statutory tariffs.

5 Conclusion

Ad valorem equivalent (AVE) tariffs are widely used as a measure of tariff protection. In this paper, we demonstrate that AVEs are not sufficient to quantify either the magnitude or sign of the effect of tariffs on economic outcomes. Put simply, AVE tariffs are not a policy variable. That is, because they depend on equilibrium prices in the presence of specific tariffs, they are not a primitive that policymakers can adjust at will. This is an especially important observation as the current US trade policy landscape, featuring tariffs at levels not seen since at least the 1930s, has produced a host of work that attempts to understand the relationship between tariffs and economic outcomes of interest in historical settings.

We show the limitations of relying on AVEs as a policy variable in the context of the years surrounding the Tokyo Round of the GATT, quantifying the effects of US tariff policy between 1972 and 1988. To do so we, construct the first dataset on annual statutory tariffs in these years and combine our data with an updated version of the Census IDB import data at a highly disaggregate level. We use these data to show that the US underwent an *accidental* liberalization between 1972 and 1979 despite the absence of wholesale policy tariff changes, as inflation combined with specific tariffs to reduce variety-level AVEs. Conversely, the negotiated tariff reductions under the Tokyo Round between 1979 and 1987 were masked in the aggregate by a shift in expenditures towards goods with higher initial tariff levels.

To quantify the effects of these changes, we extend a standard CES framework and derive a novel exact hat representation that accommodates both ad valorem and specific tariffs. This framework emphasizes that, in general, a lower AVE has ambiguous impacts on imports and aggregate price indices. In particular, the impact of falling AVEs due to rising prices depends on whether the higher prices are driven by export supply shocks, which reduce imports, or import demand shocks, which raise imports. Counterfactual exercises reveal that the inflationary erosion of specific tariff AVEs led to reductions in import price indices similar in magnitude to those driven by the statutory Tokyo Round tariff cuts. However, as the inflationary channel operated through increases in tariff-exclusive prices that our model interprets as export supply shocks, these AVE reductions were accompanied by *falling* imports. By contrast, the negotiated Tokyo Round liberalization reduced imported prices, leading to an increase in imports.

Our analysis suggests a number of avenues for future research. The 1970s and 1980s were decades characterized by dramatic changes in both trade and labor markets. We hope that the availability of annual tariff-line data encourages exploration of the relationship between trade and broader economic outcomes in these years. Further, the methods developed here allow for improved quantification of trade policy by incorporating specific tariffs. Finally, our work underscores the need for improved measurement of trade policy; in particular, we emphasize that AVE tariffs move both because of changes in statutory tariffs and also because of endogenous changes in prices, one or both of which are often ignored in empirical analysis. The importance of understanding the relationship between tariffs and economic outcomes is as high as it has been in decades. Such an understanding, however, must start with an understanding of how tariffs have changed and why.

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Supplementary Appendix for
"AVE Tariffs are a Misleading Measure of Tariff Policy"

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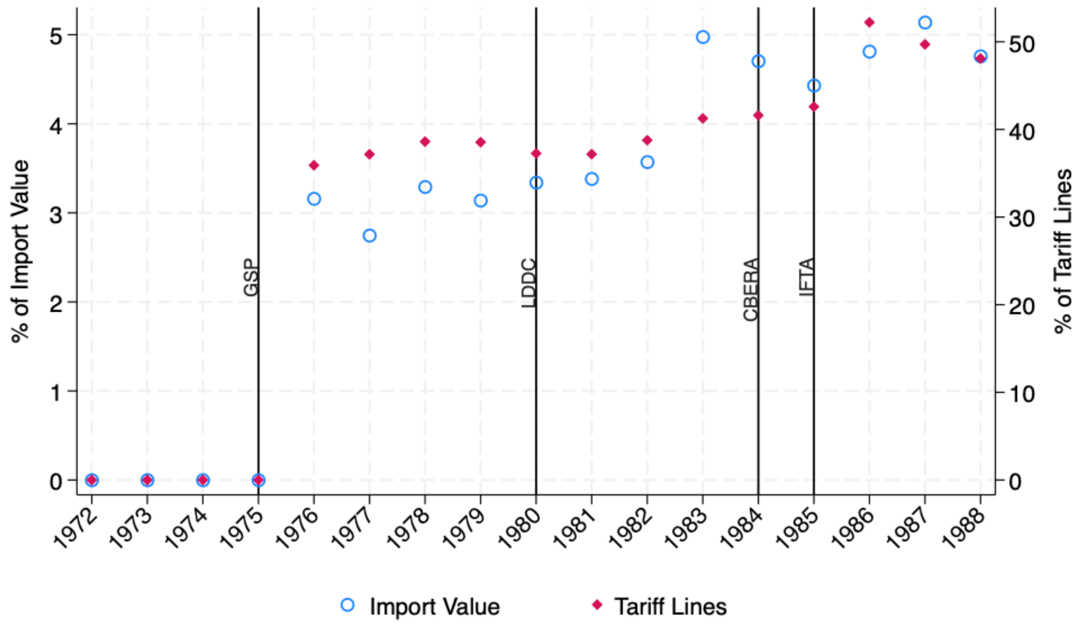
This provides an overview of the additional figures, results, and extensions of the baseline analysis found here in the appendix.

- Section A provides additional information about the timing of statutory policy changes.
  - Figure A.1 provides a time series representation of GSP, LDDC, CBERA, and IFTA tariff line shares and import shares.
  - Figure A.2 provides a decomposition of AVE tariffs into various statutory tariff rates.
- Section B details the decomposition of changes in good-level AVE tariffs into statutory and inflationary components.
- Section C details the estimation and validation of our statutory tariff database.
  - Section C.1 provides details on the statutory tariff estimation procedure and describes validation.
  - Section C.2 provides information on coverage and validation.
- Section D provides details on sectoral drivers of tariff changes.
  - Table D.1 provides sector specific changes in AVE tariffs over time.
  - Figure D.1 focuses on fuel and oil specifically.
- Section E provides full model derivations and parameter estimation details.
  - Subsection E.1 provides our baseline "hat-algebra" counterfactuals.
  - Subsection E.2 provides "exact-hat-algebra" solutions.
  - Subsection E.3 details demand and supply parameter estimation.
- Section F considers an extension of our decomposition and counterfactuals incorporating exchange rates.
- Section G documents the presence of specific tariffs in the US during the HS era from 1989-2017.

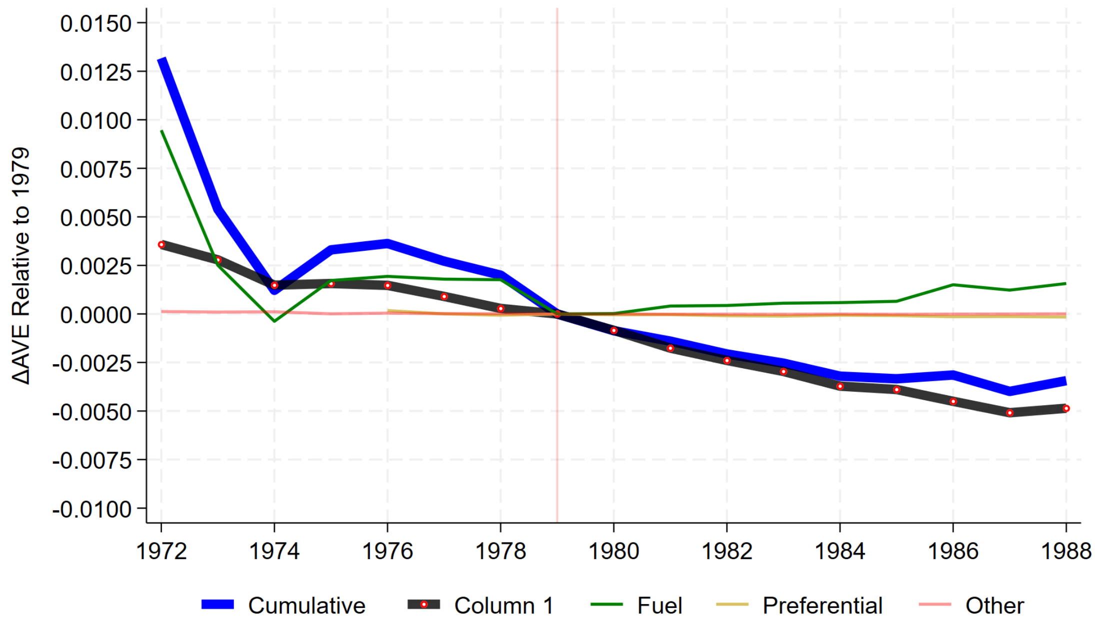
A Additional Tables & Figures

Figure A.1: Introduction of Preferential Access



Notes: Figure shows share of tariff lines covered by preferential tariff programs and the share of total US imports from country-good pairs eligible for such programs. GSP is the Generalized System of Preferences. LDDC is Least Developed Developing Countries program. CBERA is the Caribbean Basin Economic Recovery Act of 1983. IFTA is the US-Israel Free Trade Agreement.

Figure A.2: Decomposing Tariff Changes by Rate Type



Notes: Figure decomposes the AVE effect in equation (2) and Figure 2 by summing the share-weighted variety-level AVEs $s_{vgt_0}\Delta AVE_{vgt}$ into four rate type categories depending on the variety: (i) Column 1 non-fuel, (ii) Fuel (defined as five-digit TSUSA codes 475xx), (iii) Preferential tariff programs, and (iv) Other which includes column 2 non-fuel. If the US imports from an exporter under multiple rate type codes, each observation is treated as a separate "variety". See main text for details.

B Tariff Decomposition

In the main text we provide a decomposition of the changes in tariff rates between arbitrary periods t_0 and t_1. Here we detail the decomposition of intra-variety changes in AVE tariffs. Remembering that the good-level change in tariffs is simply a value-share-weighted-average (s_{vgt}) of the variety level tariffs

$$AVE_{gt} = \sum_v s_{vgt} \left(\tau_{vgt} + \frac{f_{vgt}}{p^*_{vgt}} \right) \tag{B.1}$$

fixing expenditure shares at period t_0 we can see that AVE_{gt} can vary for only three reasons. Changes in ad valorem tariff rates τ_{vgt}, specific tariff rates f_{vgt}, or prices p^*_{vgt}. Denoting the change between t_0 and period t as:

$$\begin{aligned}
\Delta AVE_t &= \left(\tau_{vgt} + \frac{f_{vgt}}{p^*_{vgt}} \right) - \left(\tau_{vgt_0} + \frac{f_{vgt_0}}{p^*_{vgt_0}} \right) \\
&= \Delta\tau_{vgt} + \frac{f_{vgt_0}}{p^*_{vgt}} + \frac{\Delta f_{vgt}}{p^*_{vgt}} - \frac{f_{vgt_0}}{p^*_{vgt_0}} \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} \frac{p^*_{vgt_0}}{p^*_{vgt}} + \frac{f_{vgt_0}}{p^*_{vgt}} - \frac{f_{vgt_0}}{p^*_{vgt_0}} \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} \left(\frac{p^*_{vgt} - \Delta p^*_{vgt}}{p^*_{vgt}} \right) + \frac{f_{vgt_0}}{p^*_{vgt}} - \frac{f_{vgt_0}}{p^*_{vgt_0}} \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} \left(1 - \frac{\Delta p^*_{vgt}}{p^*_{vgt}} \right) + \frac{f_{vgt_0}}{p^*_{vgt}} - \frac{f_{vgt_0}}{p^*_{vgt_0}} \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} - \frac{\Delta f_{vgt}}{p^*_{vgt_0}} \frac{\Delta p^*_{vgt}}{p^*_{vgt}} + f_{vgt_0} \left(\frac{p^*_{vgt_0} - p_{vgt}}{p^*_{vgt_0} p_{vgt}} \right) \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} - \frac{\Delta f_{vgt}}{p^*_{vgt_0}} \frac{\Delta p^*_{vgt}}{p^*_{vgt}} + f_{vgt_0} \left(\frac{p^*_{vgt_0} - p_{vgt}}{p^*_{vgt_0} p_{vgt}} \right) \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} - \frac{\Delta f_{vgt}}{p^*_{vgt_0}} \frac{\Delta p^*_{vgt}}{p^*_{vgt}} - \frac{f_{vgt_0}}{p^*_{vgt_0}} \frac{\Delta p^*_{vgt}}{p^*_{vgt}} \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} - \frac{\Delta p^*_{vgt}}{p^*_{vgt}} \left(\frac{\Delta f_{vgt}}{p^*_{vgt_0}} + \frac{f_{vgt_0}}{p^*_{vgt_0}} \right) \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} - \frac{\Delta p^*_{vgt}}{p^*_{vgt}} \frac{f_{vgt}}{p^*_{vgt_0}} \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} - \frac{\Delta p^*_{vgt}}{p^*_{vgt_0}} \frac{f_{vgt}}{p^*_{vgt}}
\end{aligned}$$

$$
\begin{aligned}
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} - \frac{\Delta p^*_{vgt}}{p^*_{vgt_0}} \frac{f_{vgt}}{p^*_{vgt}} \left(\frac{\tau_{vgt} + f_{vgt}/p^*_{vgt}}{\tau_{vgt} + f_{vgt}/p^*_{vgt}} \right) \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} - \frac{\Delta p^*_{vgt}}{p^*_{vgt_0}} \frac{f_{vgt}}{p^*_{vgt}\tau_{vgt} + f_{vgt}} \left(\tau_{vgt} + f_{vgt}/p^*_{vgt} \right) \\
&= \Delta\tau_{vgt} + \frac{\Delta f_{vgt}}{p^*_{vgt_0}} - \frac{\Delta p^*_{vgt}}{p^*_{vgt_0}} STS_{vgt} AVE_{vgt}.
\end{aligned}
$$

C Data Appendix

C.1 Estimating Tariffs 1972-1988

Our procedure utilizes country-good-year variation in prices and quantities to recover statutory ad valorem and specific tariffs. In doing so, we directly handle situations where the data record imports in multiple units of quantity and take care to provide tariff estimates only on the unit of quantity subject to duties. Further, we do not impose any common unit value assumptions regarding goods from different countries subject to the same statutory tariff. This is important as it is well known that additive trade barriers distort the relative prices of goods Alchian and Allen (1964). In fact, we rely on different varieties of a good having different unit values in our estimation procedure. We detail all of these issues, external validation, and corrections to the underlying IDB data in this data appendix.

The import data we use to estimate tariffs are taken from the raw ASCII files from the Census IDB that underlie the data provided by Feenstra (1996). Feenstra (1996) reports the data at the country-product-year level by aggregating over (i) dutiable and non-dutiable imports and (ii) US ports. The raw data is more disaggregate: each observation represents exports by an exporting country of a TSUSA seven-digit product to a US port pair p (which consists of both a port of entry and a port of unlading) in a year t subject to tariffs defined by "rate provision code" r. These data include total duties collected as well as import values, dutiable values, and up to two measures of quantity. 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. The rate provision code indicates whether the imports are duty free under column 1 or column 2, whether the imports are duty free under a special preferential tariff program, or whether the imports are dutiable. For dutiable imports, the rate provision code also indicates whether the tariff is ad valorem, specific, or compound – i.e., a combination of ad valorem and specific duties.

Table C.1 describes the two-digit rate provision codes in the data. A first digit of value of 1, 2, 3, or 4 describes whether the tariff is duty free, specific, ad valorem, or compound, respectively. For dutiable products, a second digit of 1 or 2 indicates that the tariff is the MFN, or normal trade relations (NTR), tariff from column 1 of the US tariff schedule or the non-NTR tariff from column 2, respectively. For duty free products, a second digit of 0 indicates that imports receive a zero tariff from, as applicable, column 1 or column 2; a second non-zero digit indicates the particular preferential program that affords duty free access. The vast majority of imports fall under one of the potential combinations of these rate provision codes. For example, rate provision code 22 indicates a column 2 specific tariff, rate provision code 31 indicates a column 1 ad valorem tariff, and rate provision code 18 indicates duty free access under preferential tariff programs such as the Generalized System of Preferences or free trade agreements. The codes 5*, 7*, *3, *4 and *8 capture various infrequently used types of tariffs.

Table C.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 this appendix an exporter by seven-digit TSUSA product pair – of good g imported into port pair p under rate provision code r at time t are

$$D_{vgprt} = \tau_{grt} M_{vgprt} + f_{grt} Q_{vgprt}. \tag{C.1}$$

Here, τ_{grt} and f_{grt} are, respectively, the unknown ad valorem and specific tariffs, while M_{vgprt} and Q_{vgprt} represent import value and quantity of imports, again noting that the data in the IDB are reported at the variety-level while statutory tariffs are specified at the more aggregate five-digit level g. Note, this relationship holds exactly, even if importers have different varieties and thus different unit values.

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).[44] For five-digit product g facing rate provision code r in year t, one would calculate

$$f_{grt} = \frac{D_{vgprt}}{Q_{vgprt}} \text{ for } r = 2^* \tag{C.2}$$

$$\tau_{grt} = \frac{D_{vgprt}}{M_{vgprt}} \text{ for } r = 3^* \tag{C.3}$$

$$D_{vgprt} = \tau_{grt} M_{vgprt} + f_{grt} Q_{vgprt} \text{ for } r = 4^*. \tag{C.4}$$

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

For specific and ad valorem tariffs, only a single variety-port-level observation would be needed using (C.2) and (C.3). But two variety-port-level observations would be needed to calculate τ_{grt} and f_{grt} for compound tariffs, as (C.4) is a single equation with two unknowns.[45]

However, multiple forms of measurement error make direct calculation infeasible. First, sometimes the unit of quantity used in the IDB and the TSUSA are different. Second, sometimes the unit of quantity varies within a tariff-line. Third, database entry error sometimes occurs. For example, as discussed in Section 2.1, the rate provision code or unit of quantity are sometimes recorded incorrectly. Additionally, years in which countries transition from column 2 to column 1 tariffs have observations where the IDB rate provision code is inconsistent with the IDB calculated duties. Fourth, the IDB data round import values and quantities to the nearest integer.

An additional complication is that many products with specific tariffs report two distinct units of quantity and associated quantity values for a given observation. Without information on which unit the IDB uses for calculated duties, there are two possible values for the specific tariff.[46] In such cases, one needs two variety-port observations to calculate the correct f_{grt}. One would then calculate a value of f_{grt} using (C.3) for each of the two units of quantity and identify the correct f_{grt} as the one that is constant across the two varieties.[47] Similarly, one would need two sets of variety-port 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 (C.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-port pairs. These would be the correct τ_{grt} and f_{grt}.

Ultimately, we can only estimate the five-digit column 1 tariffs τ_{grt} and f_{grt} with noise. However, Section C.2 shows that the resulting estimation error is negligible for the vast majority of imported goods. Our estimation strategy employs a three-step iterative process. The first step provides initial estimates of the annual tariffs. The second step refines the

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

[46]For example, TSUSA product 37315 covers "non-ornamented, wool neckties for men and boys". In 1978, this had a compound tariff of 10.5% ad valorem tariff 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.

[47]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.

estimates by removing observations that exhibit inconsistencies plausibly related to the measurement error issues 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 in which the variation in the good's tariff is sufficiently small. This final step allows us to smooth estimation error at the annual frequency that would otherwise lead us to mistakenly infer minor fluctuations in statutory tariffs.

In step one, we estimate τ_{grt} and f_{grt} from equation (C.1) separately by product g, rate provision code r, and year t between 1974 and 1988 using the following estimating equation:

$$D_{vgrt} = \tau_{grt} M_{vgprt} + f_{grt} Q_{vgprt} + \epsilon_{vgprt}. \tag{C.5}$$

We impose $\tau_{grt} = 0$ for goods facing only specific tariffs and $f_{grt} = 0$ for goods facing only ad valorem tariffs (based on their rate provision code r). If two units of quantity are provided, we estimate equation (C.5) separately for each quantity. If estimated tariffs under each of these quantities provide non-negative estimated tariffs, we use the estimates that yield the higher R^2.[48][49]

After obtaining initial tariff estimates $\hat{\tau}_{grt}$ and $\hat{f}_{grt}$ for all five-digit product by rate provision code pairs, we proceed to the second estimation step. Measurement error for individual observations is occasionally quite large due to the measurement error issues discussed above. We can thus improve the accuracy of our tariff estimates by iteratively removing observations with substantial measurement error.

To this end, we define AVE tariffs observed in the data and our estimates of them, respectively, as:

$$AVE_{vgprt} \equiv \frac{D_{vgprt}}{M_{vgprt}} \text{ and } \widehat{AVE}_{vgprt} \equiv \hat{\tau}_{grt} + \frac{\hat{f}_{grt}}{p^*_{vgprt}} \tag{C.6}$$

[48] We thank an anonymous reader for recognizing that we had incorrectly written that we use tariff rate with the lower R^2. That was a typo.

[49] In some infrequent cases, goods receive distinct specific tariffs on different components of the individual good. Because the Census IDB measures at most two distinct quantities, neither digitization nor estimation can recover the tariff on goods with with distinct specific tariffs on more than two units of quantity. Due to the infrequency of such tariff lines a concern of over-fitting, we do not estimate an expanded version of equation (C.5) for goods that have more than a single specific tariff.

where $p^*_{vgrt} = \frac{M_{vgrt}}{Q_{vgrt}}$ is the unit value. For small AVEs, the log estimation gap

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

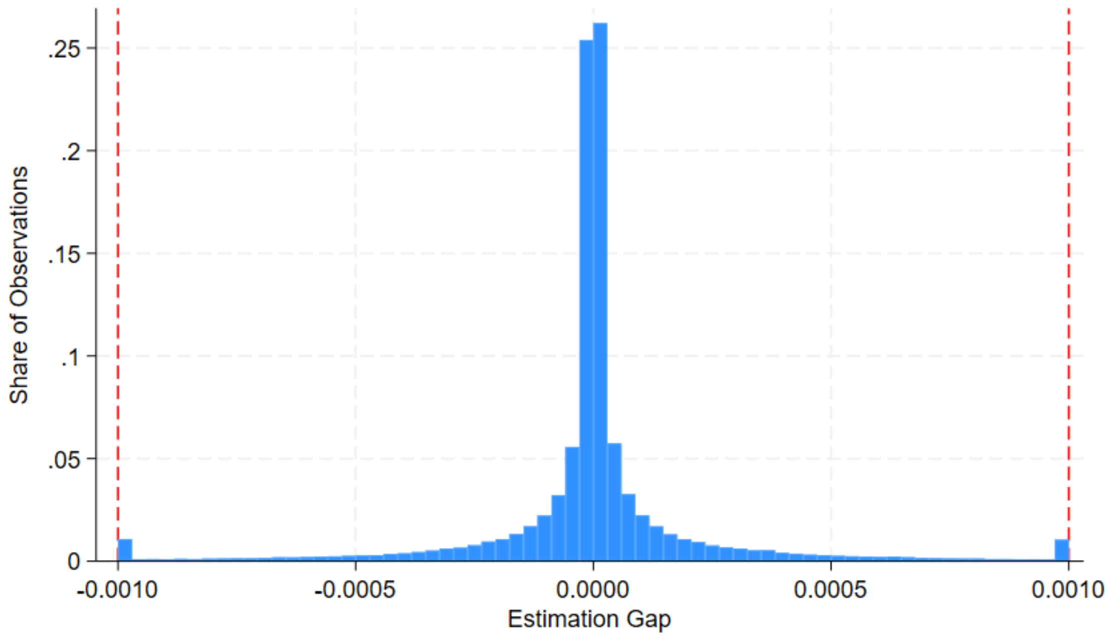
approximates the percentage difference between the tariff-inclusive price in the raw data and the value implied by our estimates of the underlying statutory tariffs. In this step we remove variety-port (vp) observations with the largest estimation gap within each cluster of good-rate-provision-year (grt) observations. For products in which the mean absolute gap $|\varepsilon^{Gap}_{vgprt}|$ exceeds 0.001, approximately 0.1% of the tariff inclusive price, we identify the single observation with the largest standardized gap and remove it from our estimating sample.

Using the reduced sample, we repeat the above two steps, again estimating tariffs and removing observations with the largest estimation gap within a cluster. We repeat this iterative process ten times. After this step, approximately 10% of dutiable observations lie in good-rate-provision-year (grt) clusters with a mean absolute estimation gap $|\varepsilon^{Gap}_{vgprt}|$ above 0.001. We manually check these observations in the TSUSA and update our dataset accordingly.

The roughly 90% of remaining dutiable observations fall in good-rate-provision-year (grt) clusters with a mean absolute estimation gap $|\varepsilon^{Gap}_{vgprt}|$ below 0.001. Figure C.1 plots the associated distribution of $\varepsilon^{Gap}_{vgprt}$, showing a near-degenerate distribution at zero that tapers symmetrically around zero. The spikes at -.001 and .001 are hard-coded observations with, respectively, $\varepsilon^{Gap}_{vgprt} < -0.001$ and $\varepsilon^{Gap}_{vgprt} > 0.001$. These are observations where the value $|\varepsilon^{Gap}_{vgprt}|$ is relatively large despite the cluster grt having a mean absolute estimation gap $|\varepsilon^{Est}_{vgprt}|$ below .001. We interpret these infrequent observations as errors in the IDB data as they provide information on duties and imports that is notably inconsistent with the information provided by other observations in the same grt cluster.

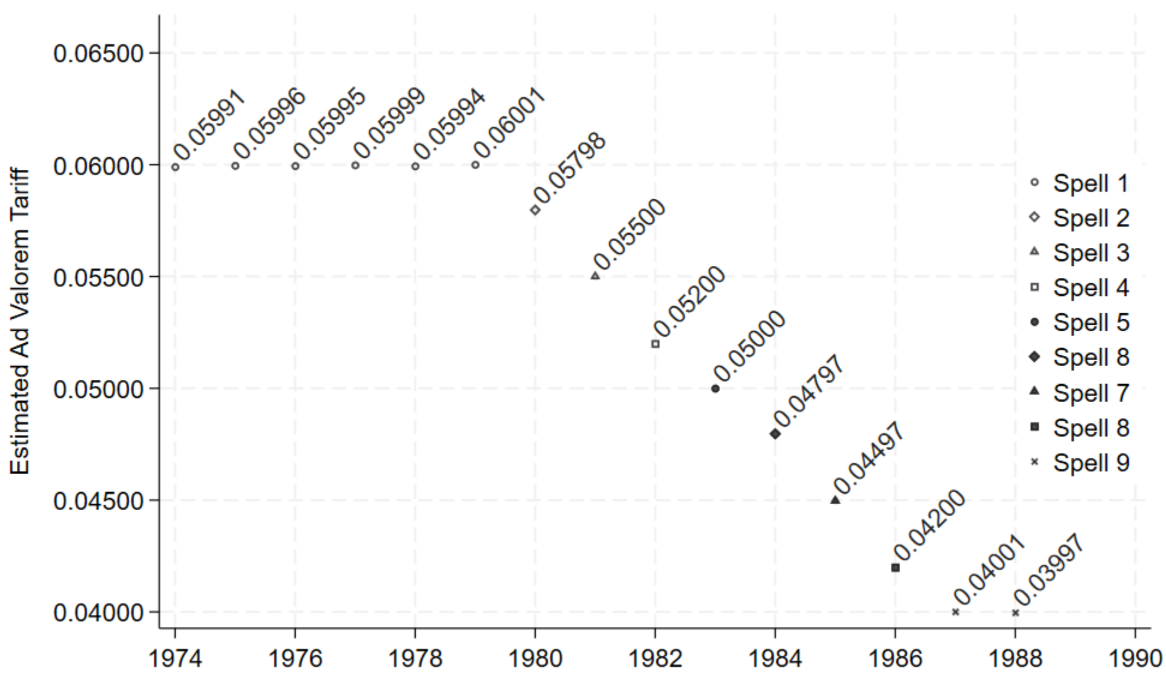
Due to estimation error, our annual estimates often exhibit minor fluctuations around a common mean, suggesting that statutory tariffs were in fact static during this period. For example, Figure C.2 suggests that statutory tariffs for five-digit good 11137 fell annually from 1980 to 1987 but, despite minor fluctuations, were static during the years 1974-1979 and 1988-1989. Indeed, the annual TSUSA for these years verifies statutory tariffs were indeed static.

Figure C.1: Distribution of Estimation Gap



Notes: Sample of five digit-rate provision code-year (grt) observations with mean absolute estimation gap $|\varepsilon^{Gap}_{vgprt}| < 0.001$ after our process of iteratively removing observations within grt clusters that have largest standardized error and re-estimating. Estimation gap defined in equation (C.7) at exporter-port-TSUSA seven 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 C.1. See text for further details.

Figure C.2: Identifying Spells from Annual Tariff Estimates



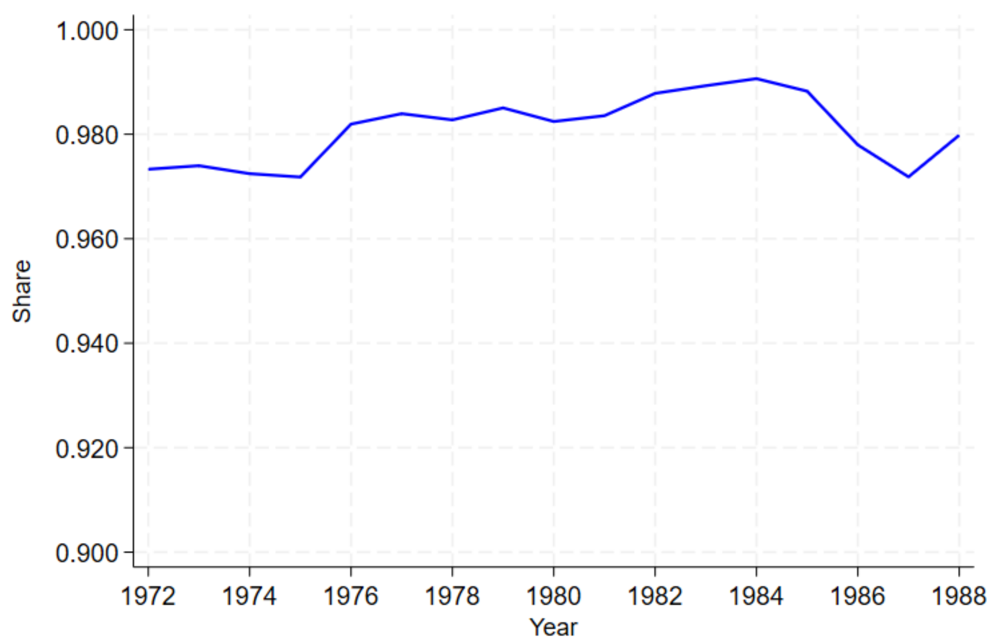
Notes: Figure displays our annual estimates of ad valorem tariffs for TSUSA 11137 "Herring, salted or pickled, but not otherwise preserved, and not in airtight containers with weight of no more than 15 lbs" before adjusting for "spells" in the data.

To rectify this issue, we begin by identifying "spells" in the data during which estimated statutory tariffs change by less than 0.015 log points in any two years. We then pool observations across time within spells and repeat steps one and two of our tariff estimation algorithm. If there are good-rate-provision-spell (grt) clusters for which observations within the cluster exhibit substantial estimation error, we manually check the TSUSA and update our dataset to reflect the true statutory tariffs.

C.2 Accuracy and Coverage of Estimated Tariffs

Our final database includes 7378 unique five-digit TSUSA goods. As shown in Figure C.3, the database covers over 97% of IDB imports annually; in each year, we do not provide estimated tariffs for 1-3% of imports. These are products which (i) we can neither estimate nor find the tariff in the TSUSA, or (ii) have "complex" tariffs that we are unable to calculate (e.g. calculating tariffs or unit values for watches and clocks in TSUSA codes 715xx requires information beyond that included in the IDB).

Figure C.3: Share of Imports Covered by Our Statutory Tariff Database



Notes: Figure displays share of IDB imports for which our database contains statutory tariffs. This includes estimated and digitized rates.

In addition to analyzing the estimation gap in Section C.1, we also use a more direct method for assessing the accuracy of our estimation strategy: we digitize 1978 statutory tariffs using the TSUSA and compare it with our estimates. The correlation is over 98% regardless of whether we look at all products, products with or without specific tariffs, column 1 or column 2 tariffs, and whether we weight by imports or not. The correlation will never be exactly 1 because (i) there is minor estimation error in our estimated tariffs (as seen in Figure C.1) and (ii) some tariff changes happen mid-year and our database is at the annual frequency. Naturally, we update our database upon identifying errors when comparing our estimates with the digitized 1978 TSUSA (when finding an error for a given five digit TSUSA code, we check the TSUSA in all other years for this good).

D Sectoral Trends Driving Aggregate AVE

Figure 2 shows the evolution of the aggregate AVE and its components neatly split into three distinct periods. To understand the sectors driving the evolution of the AVE in these three distinct periods, Table D.1 lists the sectoral breakdown of the top 25 varieties underlying the AVE and composition effects at key points in time.

The first distinct period is 1972-1975, surrounding the 1973/74 oil crisis, when the aggregate AVE drops by approximately 1 percentage point. Figure 2 shows the AVE and composition effects contribute roughly equally to the decline. Panel (a) of Table D.1 shows that "Petroleum & Natural Gas" drives the AVE effect, which is almost entirely driven by crude petroleum. Indeed, Figure D.1 shows that the dramatic fall in the fuel AVE between 1972 and 1975 is due to both (i) the temporary suspension of fuel tariffs between May 1 1973 and February 2 1975 and (ii) the five-fold increase in the price of fuel, which erodes the AVE of specific tariffs.

The oil crisis also indirectly drives the shift in the composition of US imports from higher 1979 AVE goods to lower 1979 AVE goods, as reflected by the shrinking composition effect between 1972 and 1975 in Figure 2. Indeed, Figure D.1 shows that both (i) fuel triples as a share of US imports, rising from 8% in 1972 to 25% in 1975, and (ii) fuel has a very low 1979 AVE in the midst of high fuel prices eroding the specific tariff AVE and temporary tariff suspensions during the 1979/80 oil crisis (between April 6 1979 and July 1 1980). Panel (b) of Table D.1 shows the composition shift towards fuel imports comes at the expense of, especially, textile manufacturing (e.g. "Apparel & Accessories") and "Transport Equipment" that have higher 1979 AVEs than fuel.

The second distinct time period is the 1974-1979 period, when the aggregate AVE declines by approximately 0.4%. Figure 2 shows that this is driven by the AVE effect. Panel (a) of Table D.1 shows that the 1974 values of the AVE effect are concentrated in the agricultural sectors of "Beverages", "Meats", "Tobacco", and "Vegetables". Goods in these sectors heavily rely on specific tariffs and rising 1970s global agricultural prices erode their AVE between 1974 and 1979.

The third distinct period is the 1980s, during which the aggregate AVE stays relatively constant. However, Figure 2 shows this is the result of offsetting AVE and composition effects. On the one hand, Tokyo Round tariff cuts naturally drive the AVE effect. Panel (a) of Table D.1 shows the 1988 values of the AVE effect are concentrated in manufacturing sectors

Table D.1: Distribution of AVE and Composition Effects Across Sectors

Panel (a) Share of AVE Effect Among Top 25 Exporter-by-5 Digit observations

TSUSA Part	**Δ(1972-1973)**	**Δ(1973-1974)**	**Δ(1974-1979)**	**Δ(1979-1988)**
Petroleum & Natural Gas	98.30%	81.68%		
Woven Fabrics		7.76%		
Apparel & Accessories		3.93%		
Metals, Alloys & Basic Shapes		2.35%		
Mechanical Machinery		1.01%	4.16%	5.40%
Textile Fibers		1.00%		
Beverages			22.78%	
Meats			10.93%	
Tobacco			17.04%	
Vegetables			14.20%	
Arms & Ammunition, Sporting Goods & Toys			5.36%	2.75%
Transport Equipment				19.22%
Electrical Machinery				29.16%
Optical Goods, Precision Instruments, Recording Media				11.98%
Metal Products			2.60%	7.59%
Total among top 25 obs.	98.30%	97.73%	77.07%	76.09%
Top 25 obs. share of total	85.69%	65.60%	29.15%	33.61%

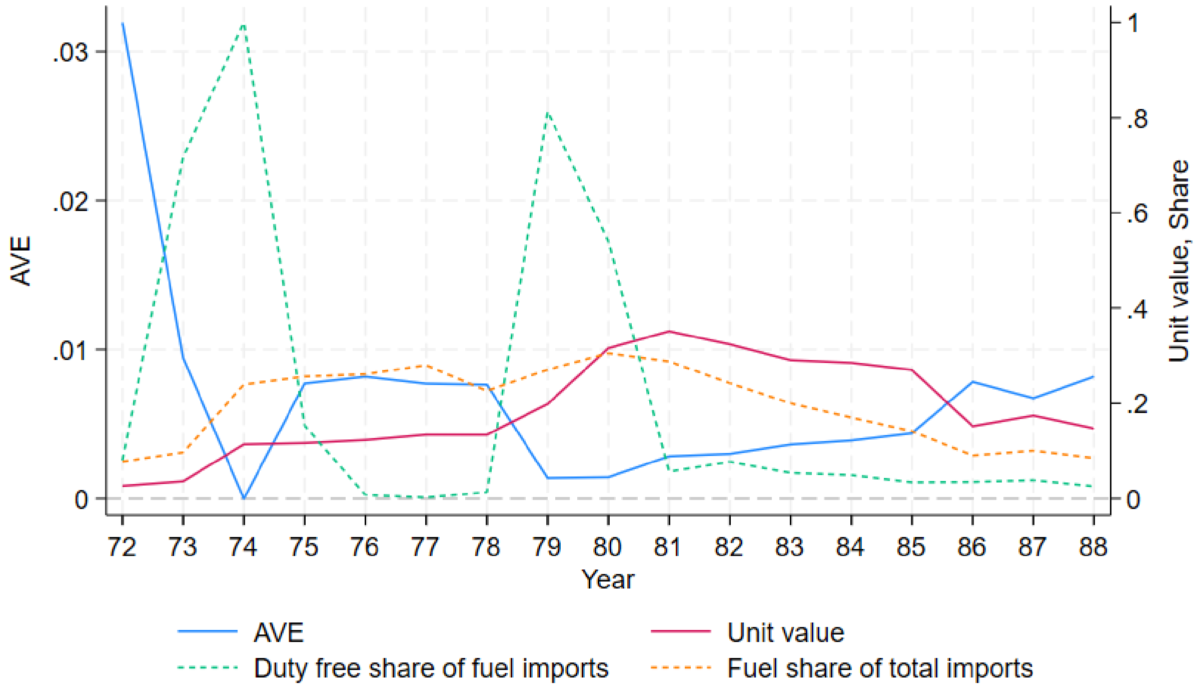
Panel (b) Share of Composition Effect Among Top 25 Exporter-by-5 Digit Observations

TSUSA Part	**Δ(1972-1973)**	**Δ(1973-1974)**	**Δ(1979-1988)**
Non-textile Manufacturing			
Transport Equipment	18.59%	12.47%	25.51%
Mechanical Machinery	7.52%	2.78%	16.03%
Electrical Machinery			14.29%
Metals, Alloys & Basic Shapes	7.68%	1.24%	
Arms & Ammunition, Sporting Goods & Toys			13.07%
Textile Manufacturing			
Apparel & Accessories	35.62%	34.47%	
Woven Fabrics	5.00%	3.85%	
Footwear, Headwear, Gloves & Bags	3.73%	5.03%	14.51%
Textile Fibers	4.29%	1.71%	
Special Fabrics	6.14%	3.77%	
Miscellaneous Textile Products			6.31%
Total among top 25 obs.	88.57%	65.32%	89.72%
Top 25 obs. share of total	138.23%	47.33%	57.28%

Notes: Table lists the TSUSA Parts with the largest changes in AVE and Composition effects from Figure 2 at key points in time using the values for the largest 25 exporter-five digit TSUSA observations. TSUSA Parts are collections of 3-digit TSUSA codes. AVE and Composition effects defined in equation (2). Because the value of any effect from equation (2) is relative to 1979, the change between year t and 1979 is equivalent to the value of the effect in year t. Sum of shares across TSUSA Parts can exceed 100% because each effect can be positive or negative.

of "Electrical Machinery", "Transport Equipment", "Optical Goods, Precision Instruments, and Recording Media", and "Metal Products". On the other hand, fuel plays an important indirect role in shifting the composition of 1980s US imports towards high 1979 AVE goods.

Figure D.1: Descriptive Statistics for Fuels



Notes: We define "fuels" as five-digit TSUSA codes 475xx. The TSUSA names these codes "Petroleum & Natural Gas". AVE defined as total duties divided by total import value. Unit value is measured in hundreds of (nominal) dollars. Unit value is the import value weighted average of variety level unit value where variety level unit value is import value divided by import quantity. AVE is measured on left y-axis, all other variables measured on right y-axis.

Indeed, Figure D.1 shows that fuel as a share of US imports falls from roughly 30% in 1979, when its AVE is close to zero, to 8% in 1988. Panel (b) of Table D.1 shows the high 1979 AVE sectors in which 1980s import expansion is concentrated. In particular, the 1988 value of the composition effect is concentrated in the textile manufacturing sectors of "Footwear, Headwear, Gloves & Bags" and "Miscellaneous Textile Products" as well as the non-textile manufacturing sectors of "Transport Equipment", "Mechanical Machinery", "Electrical Machinery", and "Arms & Ammunition, Sporting Goods & Toys". Moreover, the import growth in these sectors is driven especially by rising imports from China and Japan, both of which receive substantial tariff cuts during the 1980s. China is transitioning from column 2 to column 1 tariffs and, via the Swiss Formula, Japan is benefiting from larger Tokyo Round tariff cuts on goods with higher 1979 AVEs.

E Counterfactual Analysis

E.1 Hat Algebra

Following Fajgelbaum et al. (2020), this section uses 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 $\hat{E}_t$ and the price index for the aggregate domestic composite good $\hat{P}_{Dt}$, (ii) statutory tariffs $\hat{\tau}_{vgt}$ and $\hat{f}_{vgt}$, and (iii) structural import demand and export supply shocks $\hat{b}_{vgt}$ and $\tilde{\eta}_{vgt}$.

We make two important assumptions for analytical simplicity given the focus of our paper. First, we abstract from the production side of the economy. Second, we abstract from the impact of changes in tariff revenue on aggregate expenditure. Together, these abstractions imply that $\hat{E}_t$ and $\hat{P}_{Dt}$ are exogenous.

Utility maximizing demand at the variety level, and a variety's expenditure share of good-level expenditure, is

$$q_{vgt} = b_{vgt} p_{vgt}^{-\sigma} P_{gt}^{\sigma-1} E_{gt} \tag{E.1}$$

$$\Rightarrow s_{vgt} \equiv \frac{p_{vgt} q_{vgt}}{E_{gt}} = b_{vgt} \left(\frac{P_{gt}}{p_{vgt}} \right)^{\sigma-1} \tag{E.2}$$

where the AVE drives a wedge between tariff-inclusive and tariff exclusive prices via (7) in the main text, and the good-level price index is

$$P_{gt} = \left(\sum_{v=1}^{V_g} b_{vgt} p_{vgt}^{1-\sigma} \right)^{1/(1-\sigma)} . \tag{E.3}$$

Similarly, good-level expenditure and its share of importable-level expenditure is

$$q_{gt} = P_{gt}^{-\gamma} P_{Mt}^{\gamma-1} E_{Mt} \tag{E.4}$$

$$\Rightarrow s_{gt} \equiv \frac{E_{gt}}{E_{Mt}} = \frac{P_{gt} q_{gt}}{E_{Mt}} = \left(\frac{P_{Mt}}{P_{gt}} \right)^{\gamma-1} \tag{E.5}$$

where the importable-level price index is

$$P_{Mt} = \left(\sum_{g=1}^{G} P_{gt}^{1-\gamma} \right)^{1/(1-\gamma)}. \tag{E.6}$$

Finally, the Cobb-Douglas upper utility nest implies the aggregate price index is

$$P_t = \alpha P_{Dt} + (1-\alpha) P_{Mt} \tag{E.7}$$

with an importable expenditure share

$$\frac{E_{Mt}}{E_t} = 1 - \alpha. \tag{E.8}$$

Differentiating the price indices and using the expenditure share expressions reveals

$$\hat{P}_{gt} = \sum_{v=1}^{V_g} s_{vgt} \left(\hat{p}_{vgt} - \frac{1}{\sigma - 1} \hat{b}_{vgt} \right) \tag{E.9}$$

$$\hat{P}_{Mt} = \sum_{g=1}^{G} s_{gt} \hat{P}_{gt} \tag{E.10}$$

$$\hat{P}_t = \alpha \hat{P}_{Dt} + (1-\alpha) \hat{P}_{Mt}. \tag{E.11}$$

where the proportional change in variety-level tariff-inclusive prices is given by (18) in the main text.

To derive $\hat{p}^*_{vgt}$, we use the equilibrium tariff-exclusive price

$$\ln p^*_{vgt} = \frac{\omega}{1+\sigma\omega} \left[-\sigma \ln\left(1 + AVE_{vgt}\right) + \ln b_{vgt} + \frac{1}{\omega} \eta_{vgt} + \ln E_{gt} + (\sigma - 1) \ln P_{gt} \right] \tag{E.12}$$

where, using the expenditure shares in (E.5) and (E.8),

$$\ln E_{gt} = \ln(1-\alpha) + \ln E_t + (\gamma - 1) \ln P_{Mt} - (\gamma - 1) \ln P_{gt}. \tag{E.13}$$

Thus,

$$\hat{p}^*_{vgt} = \frac{\omega}{1+\sigma\omega} \left[-\sigma \left(1 + \widehat{AVE}_{vgt} \right) + \hat{b}_{vgt} + \frac{1}{\omega} \tilde{\eta}_{vgt} + \hat{E}_t + (\gamma - 1) \hat{P}_{Mt} + (\sigma - \gamma) \hat{P}_{gt} \right]. \tag{E.14}$$

where $\tilde{\eta}_{vgt} = d\eta_{vgt}$. Substituting (16) for the proportional change in $1 + AVE_{vgt}$ from the main text into (E.14) and rearranging yields

$$\hat{p}^*_{vgt} = \frac{\omega\varphi_{vgt}}{1+\sigma\omega}\left[-\sigma\frac{d\tau_{vgt} + \frac{df_{vgt}}{p^*_{vgt}}}{1+AVE_{vgt}} + \hat{b}_{vgt} + \frac{1}{\omega}\tilde{\eta}_{vgt} + \hat{E}_t + (\gamma - 1)\hat{P}_{Mt} + (\sigma - \gamma)\hat{P}_{gt}\right] \tag{E.15}$$

where

$$\varphi_{vgt} = \left(1 - \frac{\sigma_g\omega_g}{1+\sigma_g\omega_g}STS_{vgt}\frac{AVE_{vgt}}{1+AVE_{vgt}}\right)^{-1}. \tag{E.16}$$

Note that $\varphi_{vgt} = 1$ in the absence of specific tariffs, as $STS_{vgt} = 0$.

We now have the simultaneous equation system of endogenous variables y_t and exogenous variables x_t,

$$y_t = \left[\hat{P}_t, \hat{P}_{Mt}, \left[\hat{P}_{gt}\right]_{G\times 1}, \left[\hat{p}^*_{vgt}\right]_{V\times 1}\right]' \tag{E.17}$$

$$x_t = \left[\left[\widetilde{AVE}_{vgt}\right]_{V\times 1} \equiv \left[\frac{d\tau_{vgt} + \frac{df_{vgt}}{p^*_{vgt}}}{1+AVE_{vgt}}\right]_{V\times 1}, \left[\hat{b}_{vgt}\right]_{V\times 1}, \left[\tilde{\eta}_{vgt}\right]_{V\times 1}, \hat{P}_{Dt}, \hat{E}_t\right]', \tag{E.18}$$

governed by equations (E.9), (E.10), (E.11), and (E.15). In this system, $\hat{p}_{vgt}$ has already been eliminated from the system by substituting (18) into (E.9). 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:

$$y_t = [A_{yt} \vdots A_{xt}]\begin{bmatrix} y_t \\ x_t \end{bmatrix} \tag{E.19}$$

where

$$A_{yt} = \begin{bmatrix} 0 & s_{Mt} & 0_{1\times G} & 0_{1\times V} \\ 0 & 0 & [s_{gt}]_{1\times G} & 0_{1\times V} \\ 0_{G\times 1} & 0_{G\times 1} & 0_{G\times G} & [\beta_{0vgt}s_{vgt}]_{G\times V} \\ 0_{V\times 1} & [\beta_{2vgt}]_{V\times 1} & [\beta_{3vgt}]_{V\times G} & 0_{V\times V} \end{bmatrix} \tag{E.20}$$

(E.21)

and

$$A_{xt}=\left[\begin{array}{ccccc} 0_{1\times V} & 0_{1\times V} & 0_{1\times V} & s_{Dt} & 0\\ 0_{1\times V} & 0_{1\times V} & 0_{1\times V} & 0 & 0\\ \left[s_{vgt}\right]_{G\times V} & \left[-\frac{s_{vgt}}{\sigma-1}\right]_{G\times V} & 0_{G\times V} & 0_{G\times 1} & 0_{G\times 1}\\ \left[\beta_{4vgt}\right]_{V\times V} & \left[\beta_{5vgt}\right]_{V\times V} & \left[\beta_{6vgt}\right]_{V\times V} & 0_{V\times 1} & \left[\beta_{5vgt}\right]_{V\times 1} \end{array}\right] \tag{E.22}$$

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

Solving equation (E.19),

$$y_t = \left(I - A_{yt}\right)^{-1} A_{xt}x_t \tag{E.23}$$

which defines the endogenous variables y_t in terms of the exogenous variables x_t and the parameters in A_{xt} and A_{yt}. Even though the system (E.23) uses very large matrices with our data and requires matrix inversion, MATLAB solves the system with relatively little computational burden.

In the special case of $\omega = 0$, the system simplifies greatly. Using (18), (E.9), and (E.12):

$$\hat{p}^*_{vgt} = \tilde{\eta}_{vgt} \tag{E.24}$$

$$\Rightarrow \hat{P}_{gt} = \sum_{v=1}^{V_g} s_{vgt}\left[\frac{d\tau_{gt} + df_{gt}\frac{1}{p^*_{vgt}}}{1+AVE_{vgt}} + \left(1 - STS_{vgt}\frac{AVE_{vgt}}{1+AVE_{vgt}}\right)\tilde{\eta}_{vgt} - \frac{1}{\sigma-1}\hat{b}_{vgt}\right]. \tag{E.25}$$

Together with (E.10) and (E.11), the endogeneous $\hat{p}^*_{vgt}, \hat{P}_{gt}, \hat{P}_{Mt}$ and $\hat{P}_t$ can thus be written directly as initial share-weighted averages of the exogenous variables $\tilde{\eta}_{vgt}, \hat{b}_{vgt}, d\tau_{vgt}$ and df_{vgt}.

E.2 Exact Hat Algebra

For our exact hat algebra analysis, we redefine $\hat{x}$ so that $\hat{x} \equiv \frac{x'}{x}$ where x' is the counterfactual value and x is the initial value of a variable z.

We start by deriving the exact hat form of the price indices. Using the formula for a Cobb-Douglas price index and well-known results in the literature due to Sato (1976) and

Vartia (1976), we have

$$\hat{P}_t = \hat{P}_{Mt}^{1-\alpha}\hat{P}_{Dt}^{\alpha} \tag{E.26}$$

$$\hat{P}_{Mt} = \prod_{g\in G}\hat{P}_{gt}^{w_{gt}} \tag{E.27}$$

$$\hat{P}_{gt} = \prod_{v\in V_g}\hat{p}_{vgt}^{w_{vgt}} \tag{E.28}$$

where the exact weights are

$$w_{vgt} = \frac{s_{vgt}\left(\hat{s}_{vgt}-1\right)/\ln\hat{s}_{vgt}}{\sum_g s_{vgt}\left(\hat{s}_{vgt}-1\right)/\ln\hat{s}_{vgt}} \tag{E.29}$$

$$w_{gt} = \frac{s_{gt}\left(\hat{s}_{gt}-1\right)/\ln\hat{s}_{gt}}{\sum_g s_{gt}\left(\hat{s}_{gt}-1\right)/\ln\hat{s}_{gt}} \tag{E.30}$$

and the exact hat tariff-inclusive expenditure shares are

$$\hat{s}_{vgt} = \frac{(1+\widehat{AVE}_{vgt})^{1-\sigma}\hat{p}_{vgt}^{*1-\sigma}}{\sum_{v=1}^{V_g} s_{vgt}(1+\widehat{AVE}_{vgt})^{1-\sigma}\hat{p}_{vgt}^{*1-\sigma}} \tag{E.31}$$

$$\hat{s}_{gt} = \frac{\hat{P}_{gt}^{1-\gamma}}{\sum_{g=1}^{G} s_{gt}\hat{P}_{gt}^{1-\gamma}}. \tag{E.32}$$

Using (7), the exact hat variety-level tariff-inclusive price is

$$\hat{p}_{vgt} = \hat{p}_{vgt}^{*}(1+\widehat{AVE}_{vgt}). \tag{E.33}$$

For the exact hat tariff-exclusive variety-level price $\hat{p}_{vgt}^{*}$, we first derive the exact hat good-level equilibrium expediture. Using (E.5) and (E.8),

$$\hat{E}_{gt} = \hat{E}_t\hat{P}_{Mt}^{\gamma-1}\hat{P}_{gt}^{1-\gamma}. \tag{E.34}$$

Thus, using (E.34) and (E.12),

$$\hat{p}_{vgt}^{*} = \left[(1+\widehat{AVE}_{vgt})^{-\sigma}\hat{E}_t\hat{P}_{Mt}^{\gamma-1}\hat{P}_{gt}^{\sigma_g-\gamma}\hat{b}_{vgt}(\widehat{\exp(\eta_{vgt})})^{\frac{1}{\omega}}\right]^{\frac{\omega}{1+\sigma\omega}}, \tag{E.35}$$

where $(1+\widehat{AVE}_{vgt})$ itself depends on $\hat{p}^*_{vgt}$:

$$(1+\widehat{AVE}_{vgt}) = (\widehat{1+\tau_{gt}})\left(1 - STS_{vgt}\frac{AVE_{vgt}}{1+AVE_{vgt}}\right) + \frac{\hat{f}_{gt}}{\hat{p}^*_{vgt}} STS_{vgt}\frac{AVE_{vgt}}{1+AVE_{vgt}}. \tag{E.36}$$

We now have a 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 (E.33), the tariff-exclusive prices $\left[\hat{p}^*_{vgt}\right]_{V\times 1}$ characterized by (E.26)-(E.28) and (E.35), (ii) using (E.29)-(E.30), the expenditure shares $[\hat{s}_{gt}]_{G\times 1}$ and $[\hat{s}_{vgt}]_{V\times 1}$ characterized by equations (E.31)-(E.32), and (iii) the AVEs $\left[(1+\widehat{AVE}_{vgt})\right]_{V\times 1}$ characterized by equation (E.36). These endogenous variables depend on the following exogenous variables: (i) initial expenditure shares $[s_{gt}]_{G\times 1}$, and $[s_{vgt}]_{V\times 1}$, (ii) initial values $[AVE_{vgt}]_{V\times 1}$ and $[STS_{vgt}]_{V\times 1}$, (iii) changes in statutory tariffs $\left[(\widehat{1+\tau_{gt}})\right]_{V\times 1}$ and $\left[\hat{f}_{gt}\right]_{V\times 1}$, (iv) changes in aggregate expenditure $\hat{E}_t$ and the domestic price index $\hat{P}_{Dt}$, and (v) import demand shocks $\left[\hat{b}_{vgt}\right]_{V\times 1}$ and export supply shocks $\left[(\widehat{\exp(\eta_{vgt})})\right]_{V\times 1}$.

The exact hat algebra system is substantially more complicated to solve than the hat algebra system. First, it has roughly twice as many endogenous variables as the hat algebra system, which was already a high-dimensional system. The exact hat algebra system not only has endogenous prices $\hat{P}_{gt}$ and $\hat{p}^*_{vgt}$ from the hat algebra system, but also endogenous shares $\hat{s}_{gt}$ and $\hat{s}_{vgt}$. While the exact hat algebra change in price indices and prices requires knowing the counterfactual change in expenditure shares, the hat algebra change in price indices and prices only requires knowing the initial expenditure shares because it is a first order approximation. Second, specific tariffs add an additional layer of non-linearity to the exact hat algebra system. Equation (E.35)-(E.36) imply $\hat{p}^*_{vgt}$ enters the right hand side of (E.35) only in the presence of specific tariffs. Moreover, it does so non-linearly. In the hat algebra system, $\hat{p}^*_{vgt}$ also enters the right hand side of (E.12) only in the presence of specific tariffs, but it does so linearly via (16) because it is a first order approximation.

That said, as with the hat algebra system, the exact hat algebra system simplifies dramatically when $\omega = 0$. To begin, (E.35) reduces to

$$\hat{p}^*_{vgt} = \widehat{\exp(\eta_{vgt})} \tag{E.37}$$

so that the right hand side of (E.35) no longer depends on $\hat{p}^*_{vgt}$. This removes an important non-linearity from the exact hat algebra system. Moreover, $\hat{p}^*_{vgt}$ now only depends on the single exogenous variable related to export supply shocks. In turn, $(1+\widehat{AVE_{vgt}})$ and, hence, $\hat{p}_{vgt}$ only depend on the exogenous variables of export supply shocks and statutory tariffs via (E.33) and (E.36). In turn, the expenditure shares $\hat{s}_{vgt}$ and $\hat{s}_{gt}$ and the weights w_{vgt} and w_{gt} also only depend on these same exogenous variables via (E.29)-(E.32). The same is true for the price indices $\hat{P}_t$, $\hat{P}_{gt}$ and $\hat{P}_{vgt}$ via (E.26)-(E.28). Thus, one can simply calculate the endogenous variables of the exact hat algebra system if one has data for these exogenous shocks as well as data on the initial expenditure shares s_{gt} and s_{vgt} as well as the initial values for AVE_{vgt} and STS_{vgt}.

E.3 Estimating Elasticities

Counterfactual estimation requires elasticity estimates. For this, we use equations (10)-(14) in the main text and the approach developed by Zoutman et al. (2018) and used in a trade context by Fajgelbaum et al. (2020). Zoutman et al. (2018) show that one can recover supply and demand elasticities using exogenous changes in a tax as a single instrument. Intuitively, AVE changes trace out the wedge between tariff-inclusive and tariff-exclusive prices. As a result, the relationship between the tariff-inclusive price and the equilibrium quantity reveals the import demand elasticity, while the relationship between the tariff-exclusive price and the equilibrium quantity reveals the export supply elasticity.

Following Fajgelbaum et al. (2020), we first-difference equations (10) and (11) and place a fixed effect structure on the unobserved terms $\varphi_{\cdot,it} + u_{\cdot,vgt}$. This yields the estimating equations

$$\Delta \ln q_{vgt} = -\sigma \Delta \ln p_{vgt} + \gamma_{it} + \gamma_{vs} + \gamma_{vt} + \varepsilon_{vgt} \tag{E.38}$$

$$\Delta \ln p^*_{vgt} = \omega \Delta \ln q_{vgt} + \gamma_{it} + \gamma_{vs} + \gamma_{vt} + \varepsilon_{vgt}, \tag{E.39}$$

in which we use tariff changes to instrument for tariff-inclusive price changes in equation (E.38) and for import quantity changes in equation (E.39). The γ terms denote fixed effects: γ_{it} controls for the industry-level expenditure and price indices, γ_{vs} controls for factors such as exporter sector-specific comparative advantage, and γ_{vt} accounts for factors such as bilateral exchange rate movements. Given the magnitude of non-tariff price changes during our sample, this rich set of fixed effects helps isolate variation in prices and quantities related to tariffs. We use two-way clustered standard errors, clustering by exporter and good. ε_{vgt}

denotes the error terms.

Using AVE tariff changes as an instrument to recover elasticities faces two distinct challenges. First, statutory tariff reductions reflect, among other things, industry political influence. The extent of tariff liberalization thus potentially depends on both exporter- and good-level characteristics. To deal with this potential violation of the exclusion restriction, we rely on institutional details of the Tokyo Round negotiations and the US tariff code. As noted above, the Swiss Formula served as the starting point for Tokyo Round tariff cuts. The formula compressed the distribution of post-Tokyo tariffs by imposing an upper bound on final tariffs and dictating larger tariff cuts for goods with higher pre-Tokyo tariff levels. Countries generally phased in these tariff cuts over time beginning in 1980, with the US doing so linearly between 1980 and 1987. The 1987 AVE tariff for good g dictated by the Swiss Formula is

$$AVE_{g,87}^{Swiss} = \frac{z \times AVE_{g,79}}{z + AVE_{g,79}}, \tag{E.40}$$

where the parameter z represents the maximum allowed post-Tokyo tariff. Because z is the same for all goods – equal to 0.14 for the US – the only source of variation in liberalization under the Swiss Formula is differing pre-Tokyo tariffs.[50]

This, of course, raises questions regarding the determination of pre-Tokyo tariffs. In particular, prior GATT rounds may have produced non-random 1979 statutory tariffs. To address this, we use 1979 column 2 tariffs in the Swiss Formula rather than MFN tariffs. Column 2 tariffs were determined largely by the "Smoot-Hawley" tariffs embodied in the Tariff Act of 1930 and reflect the US MFN tariffs prior to liberalization under the Reciprocal Trade Agreements Act of 1934 and subsequent GATT rounds. As such, 1979 column 2 tariffs are highly correlated with their 1979 MFN counterparts, but are less subject to the political economy concerns outlined above because they were set 50 years prior to the Tokyo Round.[51]

Even if statutory column 2 tariffs are exogenous, a second challenge is that the 1979 AVE of column 2 specific tariffs depends on 1979 tariff-exclusive prices. This will lead to a violation of the exclusion restriction if these prices are correlated with subsequent import demand and export supply shocks. For example, all else equal, industries with stronger 1970s global supply growth will tend to have lower 1979 tariff-exclusive prices. Our estimates will be biased if the supply shocks driving the 1970s expansion persist into the 1980s.

[50] The tariff cut implied by the Swiss formula tariff is $AVE_{g,87}^{Swiss} - AVE_{g,79} = -\frac{AVE_{g,79}^2}{z+AVE_{g,79}}$.

[51] By 1979, column 2 tariffs were only applied on imports from some communist countries.

While the literature has long recognized potential violations of the exclusion restriction driven by endogenous statutory tariffs, violations due to the dependence of AVEs on tariff-exclusive prices in the presence of specific tariffs is not widely recognized. We take various steps to guard against such violations.[52] First, we control for specific tariff duties as a share of total variety-level duties. Second, we control directly for lagged growth in tariff-exclusive prices. Finally, we demonstrate robustness to restricting the sample to exclude goods with specific tariffs.

Formally, our instrument in equations (E.38) and (E.39) is

$$\Delta \ln \left(1 + AVE_{vgt}^{IV}\right). \tag{E.41}$$

In 1979, AVE_{vgt}^{IV} is simply the column 2 tariff $AVE_{g,79}^{col2}$. Beginning in 1987, upon completion of the Tokyo Round phase-in, AVE_{vgt}^{IV} is

$$AVE_{vgt}^{IV} = \frac{0.14 \times AVE_{g,79}^{col2}}{0.14 + AVE_{g,79}^{col2}}. \tag{E.42}$$

Between 1979 and 1987, AVE_{vgt}^{IV} follows a linear phase-in. For varieties not subject to column 1 tariffs – i.e., those facing column 2 tariffs or subject to preferential zero-tariff treatment – AVE_{vgt}^{IV} is zero throughout.

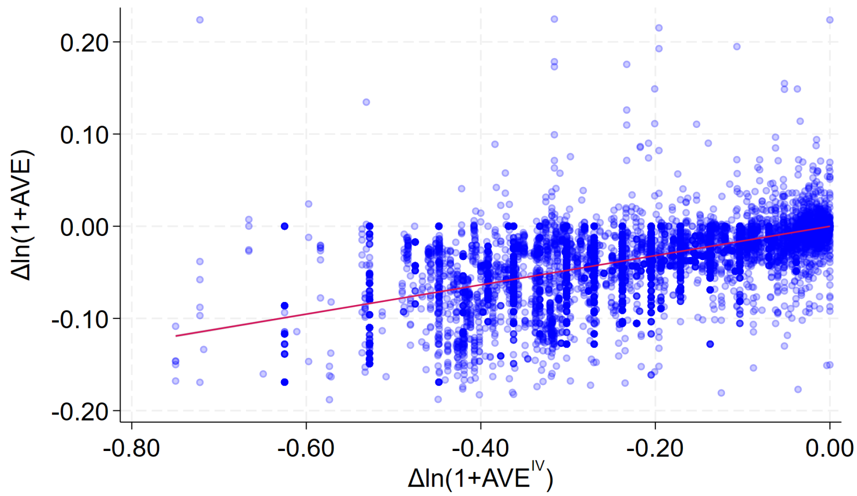
The strength of the instrument depends on the extent to which the US followed the Swiss Formula in practice. In this spirit, panel (a) of Figure E.1 plots the relationship between the Tokyo Round tariff reductions implied by our IV and observed tariff changes at the variety-year level. The clear correlation between our IV and observed AVE changes suggests that the Swiss Formula did serve as a benchmark for the implemented liberalization. Similarly, panel (b), which plots the relationship between the observed Tokyo tariff cuts and import growth between 1979 and 1987, reveals a strong relationship between the observed liberalization and import growth at the variety level.[53]

More formally, we use our instrument in Table E.1 to estimate the effect of tariff reductions on annual changes in the value of imports in panel (a) and to recover the import demand elasticity σ and inverse export supply elasticity ω in panel (b) and panel (c), respectively. Columns 1-4 in each panel employ the full sample, with column 1 reporting OLS results

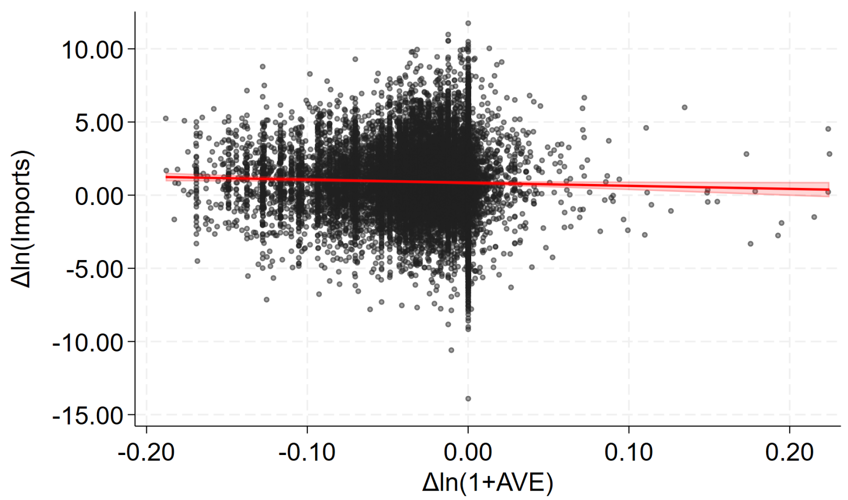
[52] We discuss these issues in detail in Greenland et al. (2026).

[53] The slope of the line of best fit in panel (b) of Figure E.1 indicates that a 1% reduction in the AVE is associated with an increase in the value of imports of approximately 2%.

Figure E.1: Imports, Tariffs, and the Swiss IV



(a) Variety-Level Tariff Changes: Observed and Instrument



(b) Variety-Level Import Growth and Tariff Changes

Notes: For the long difference between 1979 and 1987, panel (a) displays observed tariff changes $\Delta \ln (1 + AVE_{vgt})$ against instrumented tariff changes $\Delta \ln \left(1 + AVE_{vgt}^{IV}\right)$ defined by equations (E.41)-(E.42) and the surrounding text that uses the Swiss formula and column 2 tariffs. For the long difference between 1979 and 1987, panel (b) displays import growth $\Delta \ln (M_{vgt})$ against observed tariff changes $\Delta \ln (1 + AVE_{vgt})$. Point estimates are equal to 0.159 in panel (a) and -2.09 in panel (b). 95% confidence interval with standard errors clustered by exporter and five-digit TSUSA goods.

and columns 2-4 reporting IV results. To address exclusion restriction concerns related to specific tariffs, columns 3 and 4 sequentially introduce controls for specific tariff shares and lagged growth in tariff-exclusive prices.[54] As an alternative approach to dealing with these concerns, column 5 restricts the estimation sample to exclude observations with specific tariffs and repeats the IV specification from column 2. All columns include fixed effects at the exporter-year, exporter-one-digit TSUSA sector, and four-digit TSUSA good-by-year level.[55] Finally, with KP F-stats generally above the rule-of thumb value 10 and Anderson-Rubin weak instrument robust p-values generally below 0.01, our estimates appear robust against weak instrument concerns.

Panel (a) regresses log changes in import values on $\Delta \ln (1 + AVE_{vgt})$, instrumenting following equation (E.41). The coefficient is a reduced-form elasticity in that it represents the effect of tariffs on the equilibrium value of imports, but the effect depends on both the import demand and export supply elasticities. The OLS estimate of the elasticity is economically small and not close to conventional levels of statistical significance. Columns 2-5, on the other hand, report IV elasticities between -8.6 and -11.6, emphasizing the importance of addressing the endogeneity of Tokyo Round tariff cuts. Controlling for specific tariff shares as well as lagged price growth in column 4 increases the magnitude of the point estimate by approximately one-third, reflecting the importance of the exclusion restriction concerns related to specific tariffs described above. This is echoed by column 5, which restricts the sample to varieties not subject to specific tariffs and produces a similar elasticity estimate.

Panel (b) reports estimates of the import demand elasticity σ. As with the reduced-form elasticities in panel (a), the magnitude of the IV import demand elasticities is substantially larger than the OLS estimate. The IV elasticities are stable between -3.2 and -3.6 across specifications, with greater values in columns 4 and 5 that address endogeneity concerns surrounding specific tariffs. These elasticities fall well within the range found in existing work.

Finally, panel (c) reports estimates of the export supply elasticity ω. Across all columns, we find no evidence of a positively sloped export supply curve.[56] In the context of our

[54]For imports subject to column 1 tariffs, STS_{vgt} is computed using column 2 tariffs to mitigate the endogeneity concerns described above. For imports subject to column 2 tariffs or receiving duty free access under preferential tariff programs, STS_{vgt} is the observed value.

[55]Four-digit TSUSA goods differentiate among, e.g., hides and skins of bovine animals (1201x) versus other animals (1202x). Table 1 lists the one-digit TSUSA sectors.

[56]In a literal sense, the negative inverse export supply elasticities imply that the export supply curve slopes downward. This differs starkly from, e.g., Broda et al. (2008), who estimate a median inverse export supply elasticity of 1.6. That said, other papers have also found negative inverse export supply elasticities similar or larger in magnitude to our results: e.g. Basu and Fernald (1997), Antweiler and Trefler (2002), Costinot et al. (2019), and Farrokhi and Soderbery (2020).

Table E.1: Elasticity Estimates

	Panel (a): Estimates of Import Value Elasticity				
	(1)	(2)	(3)	(4)	(5)
$\Delta \ln(1 + AVE_{vgt})$	-0.473	-8.692***	-9.042***	-11.595***	-11.209***
	(0.493)	(2.511)	(2.501)	(4.396)	(3.250)
STS_{vgt}			-0.035	0.017	
			(0.043)	(0.067)	
$\Delta \ln(p^*_{vgt})$				0.216***	
				(0.041)	
1^{st} Stage Coeff.		0.137	0.133	0.114	0.149
KP F-stat		237.278	192.262	109.502	498.738
AR χ^2 p-value		0.000	0.000	0.007	0.000
	Panel (b): Estimates of σ				
	(1)	(2)	(3)	(4)	(5)
$\Delta \ln(p_{vgt})$	-1.072***	-3.280***	-3.156***	-3.461*	-3.550***
	(0.014)	(1.037)	(0.948)	(1.842)	(1.050)
STS_{vgt}			0.042	0.227	
			(0.061)	(0.145)	
$\Delta \ln(p^*_{vgt})$				-5.165	
				(4.009)	
1^{st} Stage Coeff.		0.466	0.501	0.498	0.596
KP F-stat		12.563	12.791	2.712	11.918
AR χ^2 p-value		0.000	0.000	0.002	0.000
	Panel (c): Estimates of ω				
	(1)	(2)	(3)	(4)	(5)
$\Delta \ln(q_{vgt})$	-0.366***	-0.219**	-0.238***	-0.232	-0.211***
	(0.010)	(0.089)	(0.089)	(0.152)	(0.077)
STS_{vgt}			0.022	0.074*	
			(0.020)	(0.040)	
$\Delta \ln(p^*_{vgt})$				-1.633***	
				(0.361)	
1^{st} Stage Coeff.		-1.528	-1.582	-1.722	-2.115
KP F-stat		18.083	19.240	8.643	16.007
AR χ^2 p-value		0.010	0.006	0.165	0.007
N	240,377	240,377	240,377	140,055	195,012
Specification	OLS	IV	IV	IV	τ_{vgt} Only

Notes: Dependent variable is the change in import value ΔM_{vgt} in panel (a) and, per the estimating equations (E.38)-(E.39), $\Delta \ln(q_{vgt})$ in panel (b) and $\Delta \ln(p^*_{vgt})$ in panel (c). Estimating sample period is 1979-1987. IV in columns 2-5 is the Swiss IV defined by equation (E.41). All specifications include fixed effects at the one-digit TSUSA by exporter, four-digit TSUSA by year, and exporter by year levels. Columns 1-2 and 5 include no additional controls. Columns 3-4 add, respectively, STS_{vgt} and $\Delta \ln(p^*_{vgt})$ defined as the annualized change between years t and $t-7$. Column 5 excludes observations with specific tariffs. Two-way clustered standard errors clustered by five-digit TSUSA goods and exporter.
$^*p < .1, ^{**}p < .05, ^{***}p < .01$

model, we interpret these estimates as implying a perfectly elastic export supply curve. This implies that tariff reductions are passed through entirely into lower tariff-inclusive prices, with no accompanying effect on tariff-exclusive prices. This small open economy interpretation is analogous to the results of Fajgelbaum et al. (2020), who find that US tariff increases during the first US-China trade war had no impact on tariff-exclusive prices. The result also significantly simplifies the counterfactual analysis in Section 4.

Table E.2: Elasticity Estimates with Five-Digit TSUSA by Year Fixed Effects

	Panel (a) Estimates of Import Value Elasticity				
	(1)	(2)	(3)	(4)	(5)
$\Delta\ln(AVE_{vgt})$	-0.508	-7.060**	-7.320**	-11.701*	-11.012***
	(0.549)	(3.252)	(3.311)	(5.993)	(4.026)
STS_{vgt}			-0.027	-0.045	
			(0.045)	(0.107)	
$\Delta\ln(p^*_{vgt})$				0.217***	
				(0.050)	
1^{st} Stage		0.151	0.148	0.132	0.156
KP F-stat		363.951	352.177	194.771	461.593
AR χ^2 p-value		0.029	0.027	0.052	0.006
	Panel (b) Estimates of σ				
	(1)	(2)	(3)	(4)	(5)
$\Delta\ln(p_{vgt})$	-1.069***	-2.971**	-2.802**	-4.404	-3.282***
	(0.013)	(1.253)	(1.082)	(4.632)	(1.088)
STS_{vgt}			0.055	0.475	
			(0.080)	(0.561)	
$\Delta\ln(p^*_{vgt})$				-7.484	
				(10.372)	
1^{st} Stage		0.472	0.532	0.430	0.682
KP F-stat		7.081	8.059	0.621	10.040
AR χ^2 p-value		0.007	0.006	0.054	0.001
	Panel (c) Estimates of ω				
	(1)	(2)	(3)	(4)	(5)
$\Delta\ln(q)$	-0.369***	-0.242*	-0.273**	-0.187	-0.235**
	(0.009)	(0.127)	(0.125)	(0.243)	(0.091)
STS_{vgt}			0.030	0.127***	
			(0.027)	(0.046)	
$\Delta\ln(p^*_{vgt})$				-1.803***	
				(0.577)	
1^{st} Stage		-1.402	-1.490	-1.895	-2.239
KP F-stat		7.139	7.659	3.686	10.488
AR χ^2 p-value		0.060	0.032	0.514	0.013
Obs.	236,421	236,421	236,421	136,233	191,570
Spec	OLS	IV	IV	IV	τ_{vgt} Only

Notes: Dependent variable is the change in import value ΔM_{vgt} in panel (a) and, per the estimating equations (E.38)-(E.39), $\Delta\ln(q_{vgt})$ in panel (b) and $\Delta\ln(p^*_{vgt})$ in panel (c). Estimating sample period is 1979-1987. IV in columns 2-5 is the Swiss IV defined by equation (E.41). All specifications include fixed effects at the one-digit TSUSA by exporter, five-digit TSUSA by year, and exporter by year levels. Columns 1-2 and 5 include no additional controls. Columns 3-4 add, respectively, STS_{vgt} and $\Delta\ln(p^*_{vgt})$ defined as the annualized change between years t and $t-7$. Column 5 excludes observations with specific tariffs. Two-way clustered standard errors clustered by five-digit TSUSA and country.
$^*p < .1, ^{**}p < .05, ^{***}p < .01$

We now present various robustness exercises. First, Table E.2 adjusts the analysis from

Table E.1 to include five-digit TSUSA-by-year fixed effects. In Fajgelbaum et al. (2020), the 2018-2019 trade war provides considerable variation in variety-level tariffs between targeted and non-targeted countries and products. There is less variation in our context because of the largely MFN nature of tariffs and the broad nature of the Tokyo Round liberalization. Thus, our baseline analysis uses four-digit TSUSA-by-year fixed effects. Nevertheless, Table E.2 shows that our results are very similar with five-digit TSUSA-by-year fixed effects, albeit somewhat less precise and more variable across specifications.

Table E.3: Elasticity Estimates Placebo Exercise

	Panel (a) Estimates of Import Value Elasticity			
	(1)	(2)	(3)	(4)
$\Delta\ln(AVE_{vgt})$	-0.178	-58.326	71.165	-731.890
	(0.361)	(115.526)	(73.200)	(1215.324)
STS_{vgt}			1.111	
			(1.378)	
1^{st} Stage		0.007	-0.008	0.001
KP F-stat		1.422	1.415	0.742
AR χ^2 p-value		0.489	0.346	0.446
	Panel (b) Estimates of σ			
	(1)	(2)	(3)	(4)
$\Delta\ln(p_{vgt})$	-1.056***	-2.513	-2.590	-2.802
	(0.014)	(2.446)	(1.840)	(2.491)
STS_{vgt}			-0.034	
			(0.205)	
1^{st} Stage		0.262	0.349	0.330
KP F-stat		2.044	3.310	1.707
AR χ^2 p-value		0.295	0.152	0.291
	Panel (c) Estimates of ω			
	(1)	(2)	(3)	(4)
$\Delta\ln(q_{vgt})$	-0.329***	-0.388	-0.397	-0.358
	(0.011)	(0.378)	(0.278)	(0.318)
STS_{vgt}			0.004	
			(0.090)	
1^{st} Stage		-0.659	-0.905	-0.926
KP F-stat		1.008	1.889	1.020
AR χ^2 p-value		0.148	0.052	0.171
Obs.	139,137	139,137	138,588	97,932
Spec	OLS	IV	IV	τ_{vgt} Only

Notes: Dependent variable is the change in import value ΔM_{vgt} in panel (a) and, per the estimating equations (E.38)-(E.39), $\Delta\ln(q_{vgt})$ in panel (b) and $\Delta\ln(p^*_{vgt})$ in panel (c). Estimating sample period is 1972-1979. For all years, value of the IV in columns 2-5 is the value of the Swiss IV $\Delta\ln\left(1+AVE^{IV}_{vgt}\right)$ defined by equation (E.41) for the change between 1979 and 1980. All specifications include fixed effects at the one-digit TSUSA by exporter, four-digit TSUSA by year, and exporter by year levels. Columns 1-2 and 5 include no additional controls. Columns 3-4 add, respectively, STS_{vgt} and $\Delta\ln\left(p^*_{vgt}\right)$ defined as the annualized change between years t and $t-7$. Column 5 excludes observations with specific tariffs. Two-way clustered standard errors clustered by five-digit TSUSA and country.
$^{*}p<.1,^{**}p<.05,^{***}p<.01$

Second, we address pre-trends with two alternative approaches. Table E.3 uses the pre-Tokyo period of 1972-1979 and instruments for observed AVE changes with the value of our

Swiss IV defined in equation (E.41) for the years 1979-1980. This placebo exercise does not produce the pattern of estimates found in our baseline analysis; indeed, none of the point estimates are statistically different from zero at conventional levels of significance. Finally, Table E.4 employs our baseline sample period but controls for the lagged dependent variable from seven years prior.[57] The results corresponding to panels (a) and (b) in Table E.1 change very little, while estimates of the export supply elasticity ω from panel (c) are no longer statistically significant in any IV specification.

[57] Under this approach, the observation reflecting the change between 1979 and 1980 in our post-Tokyo sample period corresponds to the lagged change between 1972 and 1973, which are the first two years in our dataset. Further, except for the changes between 1986 and 1987, all post-Tokyo sample observations correspond to lagged changes from the pre-Tokyo period.

Table E.4: Elasticity Estimates, Lagged Outcome Control

	Panel (a) Estimates of Import Value Elasticity				
	(1)	(2)	(3)	(4)	(5)
$\Delta \ln(AVE_{vgt})$	-0.856	-11.495***	-12.119***	-12.109***	-13.294***
	(0.518)	(4.380)	(4.615)	(4.606)	(4.910)
Lag $\Delta \ln(m_{vgt})$	-1.189***	-1.191***	-1.191***	-1.188***	-1.189***
	(0.033)	(0.034)	(0.034)	(0.034)	(0.036)
STS_{vgt}			-0.061	-0.062	
			(0.074)	(0.074)	
$\Delta \ln(p^*_{vgt})$				0.070*	
				(0.036)	
1^{st} Stage		0.121	0.114	0.114	0.139
KP F-stat		131.598	108.945	109.440	335.717
AR χ^2 p-value		0.006	0.007	0.007	0.004
	Panel (b) Estimates of σ				
	(1)	(2)	(3)	(4)	(5)
$\Delta \ln(p_{vgt})$	-1.010***	-3.550**	-3.242***	-3.827*	-4.122*
	(0.015)	(1.407)	(1.172)	(2.084)	(2.469)
Lag $\Delta \ln(q_{vgt})$	-0.856***	0.559	0.388	-1.348***	1.019
	(0.022)	(0.777)	(0.648)	(0.127)	(1.456)
STS_{vgt}			0.130	0.171	
			(0.105)	(0.157)	
$\Delta \ln(p^*_{vgt})$				-7.468	
				(4.666)	
1^{st} Stage		0.525	0.593	0.490	0.566
KP F-stat		4.692	5.628	2.634	2.022
AR χ^2 p-value		0.001	0.001	0.001	0.002
	Panel (c) Estimates of ω				
	(1)	(2)	(3)	(4)	(5)
$\Delta \ln(q_{vgt})$	-0.338***	-0.179	-0.232	-0.232	-0.136
	(0.010)	(0.145)	(0.152)	(0.152)	(0.175)
Lag $\Delta \ln(p^*_{vgt})$	-1.386***	-1.754***	-1.633***	-1.633***	-1.851***
	(0.040)	(0.344)	(0.361)	(0.361)	(0.406)
STS			0.074*	0.074*	
			(0.040)	(0.040)	
1^{st} Stage		-1.528	-1.582	-1.722	-2.115
KP F-stat		9.617	8.643	8.643	8.244
AR χ^2 p-value		0.246	0.165	0.165	0.467
Obs.	140,055	140,055	140,055	140,055	108,968
Spec	OLS	IV	IV	IV	τ_{vgt} Only

Notes: Dependent variable is the change in import value ΔM_{vgt} in panel (a) and, per the estimating equations (E.38)-(E.39), $\Delta \ln(q_{vgt})$ in panel (b) and $\Delta \ln(p^*_{vgt})$ in panel (c). Estimating sample period is 1979-1987. IV in columns 2-5 is the Swiss IV defined by equation (E.41). All specifications include annualized change in dependent variable between years t and $t-7$, and fixed effects at the one-digit TSUSA by exporter, four-digit TSUSA by year, and exporter by year levels. Columns 1-2 and 5 include no additional controls. Columns 3-4 add, respectively, STS_{vgt} and $\Delta \ln(p^*_{vgt})$ defined as the annualized change between years t and $t-7$. Column 5 excludes observations with specific tariffs. Two-way clustered standard errors clustered by five-digit TSUSA and country.
$^*p < .1, ^{**}p < .05, ^{***}p < .01$

F Exchange Rates and AVE

As a simple and transparent way to illustrate the potential role of exchange rates, we begin by writing the tariff-exclusive log import price as

$$\ln\left(p^*_{vgt}\right) = \ln\left(p^{*,F}_{vgt}\right) + \ln\left(XR^{\frac{\$}{F}}_{v,t}\right) \tag{F.1}$$

where $p^{*,F}_{vgt}$ is tariff-exclusive price in the currency of exporter v and $XR^{\frac{\$}{F}}_{v,t}$ is the bilateral exchange rate between the US and exporter v such that a higher $XR^{\frac{\$}{F}}_{v,t}$ represents a depreciation of the US dollar. While modeling endogenous exchange rate determination and the choice of invoicing currency is beyond the scope of our analysis, one way to interpret this exercise is to assume that exports to the US are predominantly priced in the exporter's local currency and that US tariff policy has minimal impact on observed movements in bilateral exchange rates during our sample period.

While US tariff policy could affect bilateral US exchange rates, we take the view here that the substantial observed movements in US exchange rates during our sample are driven by broader macroeconomic factors. In particular, US government bond yields peaked at more than 15% in the first half of the 1980s (Destler, 1991; Frankel, 1994, 2015), leading to a surge in global demand for US dollars and resulting trade imbalances that were met by calls for increased protection.[58] This ultimately resulted in the 1985 Plaza Accord, under which countries were encouraged to appreciate their currency relative to the dollar to help address these imbalances (Frankel, 2015). Furthermore, in contrast to a prominent literature that suggests otherwise in more modern settings (e.g. Gopinath et al. (2010)), empirical evidence supports the premise that exports to the US were largely priced in exporters' local currency during our sample period. For example, Magee et al. (1974) show that 40-80% of contracts for US imports from West Germany and Japan were denominated in the exporter's local currency. Relatedly, Hooper and Mann (1989) document that US import prices move with exchange rates in the 1980s, but that this relationship had diminished by the 1990s.[59]

To illustrate how the distinct components of USD tariff-exclusive prices in equation (F.1)

[58]See https://fred.stlouisfed.org/series/DGS10

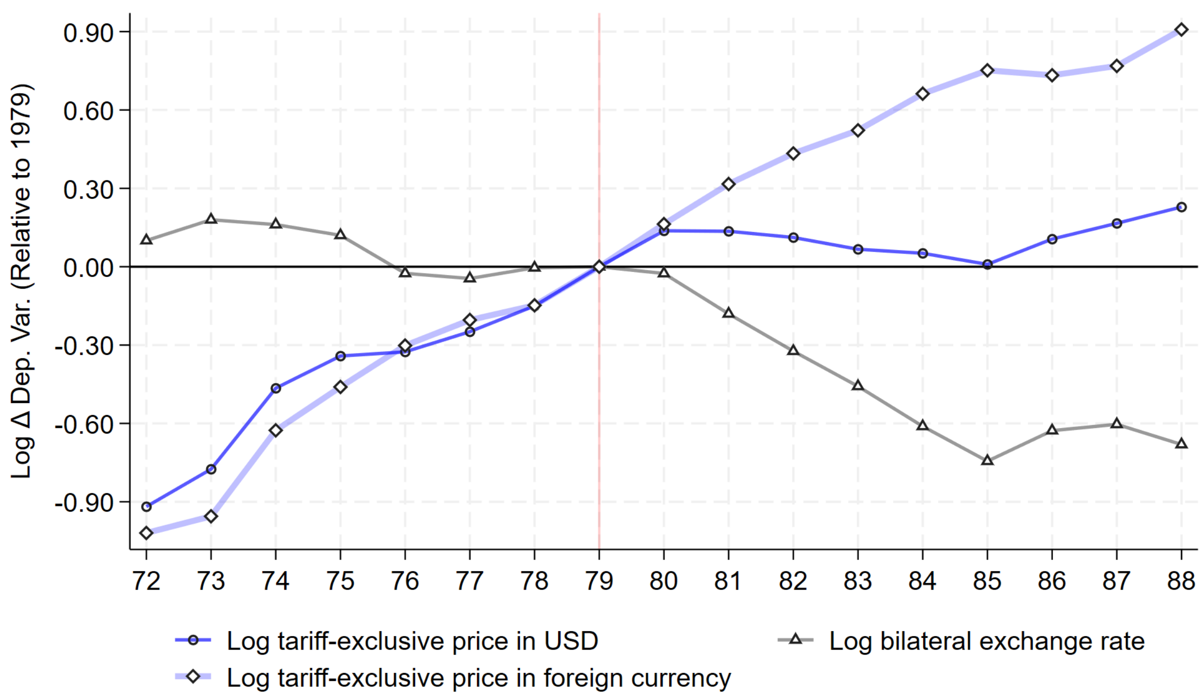
[59]See a discussion by Goldberg and Knetter (1996) for further evidence during this period.

evolve over our sample period, we use the estimating equation

$$y_{vgt} = \sum_{t'=1972, t' \neq 1979}^{1988} \beta_{y,t'} \times 1(\text{year} = t') + \gamma_v + \gamma_g + \varepsilon_{vgt}, \tag{F.2}$$

where y_{vgt} is either $\ln\left(p^*_{vgt}\right)$, $\ln\left(p^{*,F}_{vgt}\right)$, or $XR^{\frac{\$}{F}}_{v,t}$, normalized to their 1979 values.[60] The γ terms are fixed effects that remove exporter- and good-level means to facilitate comparison of the components in equation (F.1) across observations. The variable $1(\text{year} = t')$ is a year indicator variable, with 1979 the omitted year. The coefficients $\beta_{y,t'}$ are time-varying coefficients that capture average log changes in the dependent variable y relative to 1979. Finally, ε_{vgt} is the error term.

Figure F.1: Statutory and Price Driven Changes in US Column 1 Tariffs



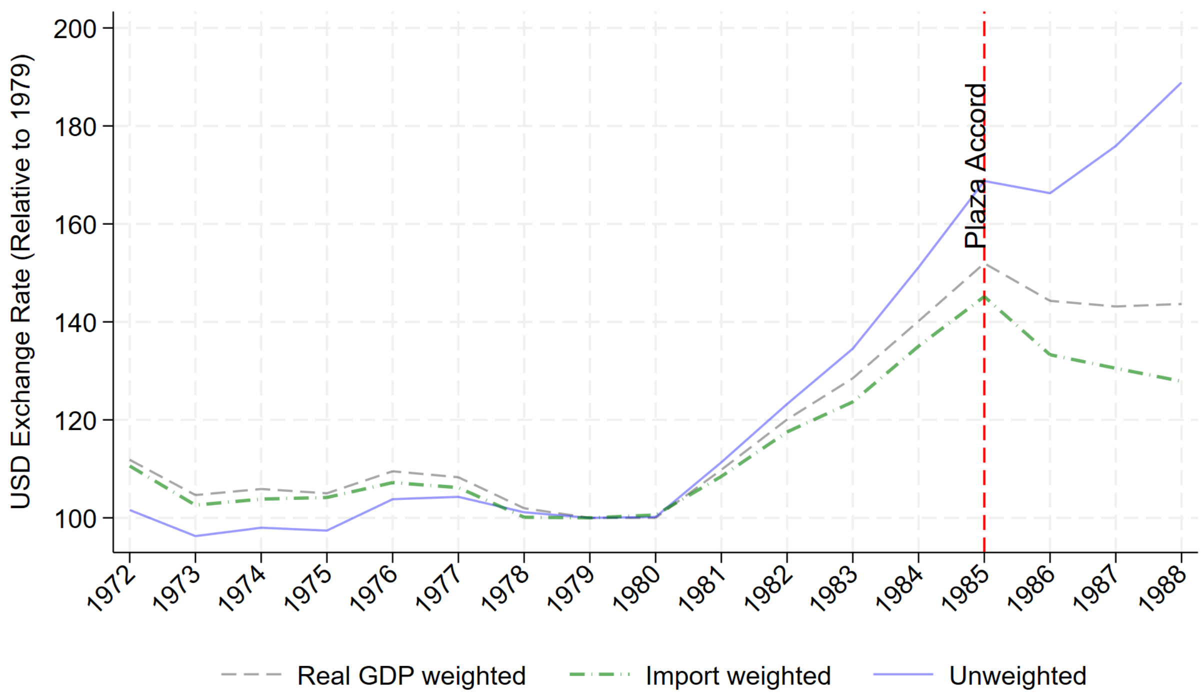
Notes: Figure displays coefficients $\beta_{y,t}$ from estimating equation (F.1) for dependent variables $y = \ln\left(p^*_{vgt}\right), \ln\left(p^{*,F}_{vgt}\right), XR^{\frac{\$}{F}}_{v,t}$. Higher (lower) $XR^{\frac{\$}{F}}_{v,t}$ represents depreciation (appreciation) of the US dollar. Regressions weighted by 1979 5-digit TSUSA import shares.

Figure F.1 shows the evolution of the coefficients $\beta_{y,t}$. While the coefficients for the USD

[60]For this exercise we focus solely on varieties subject to column 1 rates, for which statutory rates are common to all exporters. This covers 90-95% of imports annually in our sample period.

tariff-exclusive price increase between 1972 and 1979, they remain quite stable during the post-1979 period. However, this stability masks offsetting effects: the bilateral appreciation of the US dollar, especially during the early 1980s, offsets the continued increase in tariff-exclusive prices denominated in foreign currency. Indeed, Figure F.2 shows the import-weighted bilateral exchange rate appreciated by 42% between 1979 and 1985. Absent this dramatic strengthening of the US dollar, rising foreign currency tariff-exclusive prices would have continued eroding specific tariff AVEs throughout the 1980s.

Figure F.2: Evolution of USD Exchange Rate 1972-1987



Notes: Bilateral exchange rate and real GDP data from Penn World Tables version 10.01. Exchange rates indexed to 1979 at value of 100. Higher (lower) exchange rate represents appreciation (depreciation) of US dollar. Real GDP and import weighted exchange rates use country-level 1979 real GDP and imports as weights. Vertical red line at 1985 indicates the signing of the Plaza Accord under which countries were encouraged to appreciate their currency relative to the US dollar help address US trade imbalances.

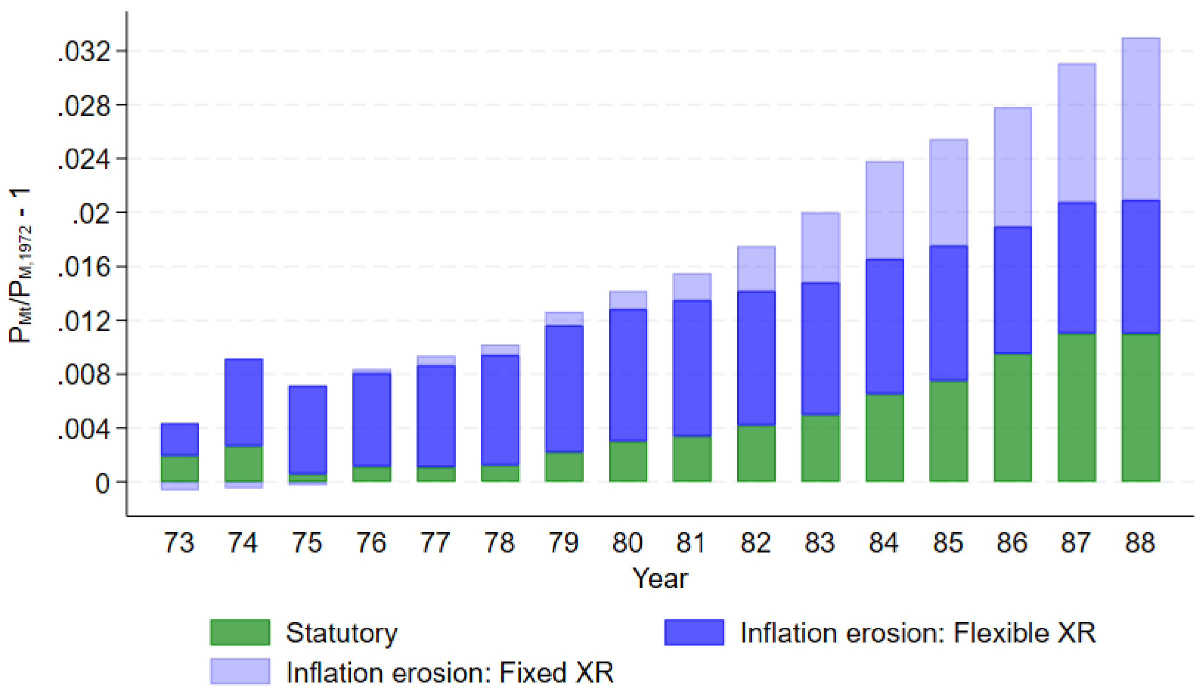
In the main text, we emphasize that AVEs move for a variety of reasons, only some of which are directly related to policy, and note that the implications for import growth depend on the particular source of AVE movements. Bilateral exchange rate movements are an additional non-policy channel through which AVEs move, and further underscore the nuanced relationship between AVEs, prices, and imports.

To illustrate, Figure F.3 revisits the first counterfactual from Section 4.1, as illustrated in

Figure 5. The figure depicts two versions of the effects related to the removal of inflationary erosion. The blue bars represent the effects from Figure 5, which remove US dollar tariff-exclusive price changes. These price changes combine both foreign currency price changes and exchange rate changes. The exercise thus reflects the removal the evolution of prices as depicted by the blue line in Figure F.1. The combined height of the blue and purple bars, on the other hand, represent the effects of removing foreign currency tariff-exclusive price changes only. These price changes hold the exchange rate component of US dollar tariff-exclusive prices fixed. The exercise thus reflects removing the evolution of prices depicted by the purple line in Figure F.1. The additional effect due to the purple bars is slightly negative in the early 1970s but grows substantially during the 1980s. By the end of the 1980s, broad appreciation of the US dollar notably limits the extent that rising tariff-exclusive prices erode specific tariff AVEs.

This observation has implications for our counterfactuals related to price shocks and the importable price index in Figure 6, and how different AVE movements affect imports in Table 2. Absent appreciation of the US dollar, the effect of price shocks on the importable price index would not have tapered off in the 1980s as in Figure 6. As a result, the observed appreciation of the US dollar represents a source of welfare gains. Additionally, by keeping tariff-exclusive US dollar prices low, appreciation of the US dollar stimulates imports while also increasing specific tariff AVEs. Thus, appreciation of the US dollar and rising tariff-exclusive foreign currency prices driven by export supply shocks each represent a separate reason for the positive relationship between imports and AVEs in column 2 of Table 2.

Figure F.3: Cumlulative Importable Price Index Increase with Fixed Exchange Rates



Notes: Figure presents exact hat counterfactual estimates of the cumulative effect of changes in AVEs on the importable price index using equation (E.27). Green (blue) bars represent the percentage increase in the importable price index if statutory tariff changes (inflation erosion of specific tariffs due to US dollar tariff-exclusive price changes) had not happened. Combined height of blue and purple bars represent percentage increase in the importable price index via higher AVEs if inflation erosion due to foreign currency tariff-exclusive price changes had not happened. Thus, the blue bar is the effect of inflation erosion that includes movements in US dollar exchange rates, and the purple bar is the additional effect due to holding US dollar exchange rates fixed.

G Specific Tariffs 1972-2017

The presence of specific tariffs complicates the relationship between AVEs and import prices. While such tariffs are prevalent in our sample period, an important question remains: are specific tariffs relevant in more recent years? In this section we show that the answer is yes, documenting their continued importance through the era of the Harmonized System (HS) and up to the eve of the 2018 trade war.

To do so, we contribute to recent improvements in the quality of US tariff data in the early HS years. Electronic versions of the US Harmonized Tariff System (USHTS) are provided by the USITC Tariff Database for 1997 onward and by Feenstra et al. (2002) for 1989-1996. However, the Feenstra et al. (2002) data omits the ad valorem tariff in cases in which the specified tariff consists of both specific and ad valorem tariffs – a "compound tariff". Recently, Acosta and Cox (2025) have provided data that rectifies this issue. We further contribute by adding the unit of quantity for specific tariffs in the 1989-1996 years, which is necessary to compute the variety-level AVE of these tariffs.[61]

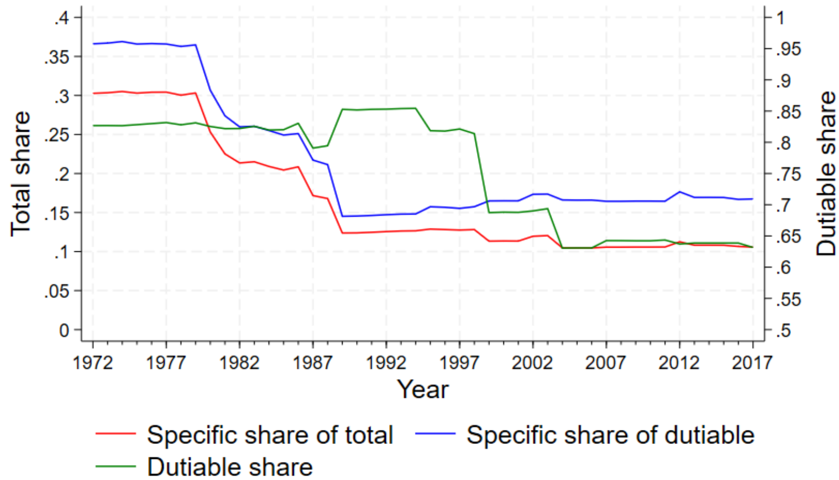
Figure G.1 depicts the evolution of the US tariff code between 1972 and 2017 for column 1 and column 2 tariffs, respectively. In the 1972-1979 pre-Tokyo years, tariff lines with specific tariffs account for approximately 30% of all column 1 lines and 37% of dutiable column 1 lines. The shares are similar, if slightly higher, for column 2 tariffs. During the Tokyo Round implementation years of 1980-1987, the share of tariff lines with specific tariffs falls considerably. By 1987, 17% of column 1 tariff lines and 22% of dutiable column 1 tariff lines have specific tariffs. While the reduction is less dramatic, specific tariffs also become less common among column 2 lines in these years.

The figure also reveals that the transition between the TSUSA and HS tariff codes in 1989 represents more than a change in classification system. For both column 1 and column 2 tariff lines, the share of dutiable lines increases notably and the share of lines with specific tariffs falls yet again. Indeed, the share of column 1 lines and column 1 dutiable lines with specific tariffs falls by roughly one-third, to 12% and 15%, respectively. The share of column 2 lines with specific tariffs falls by roughly one-quarter.

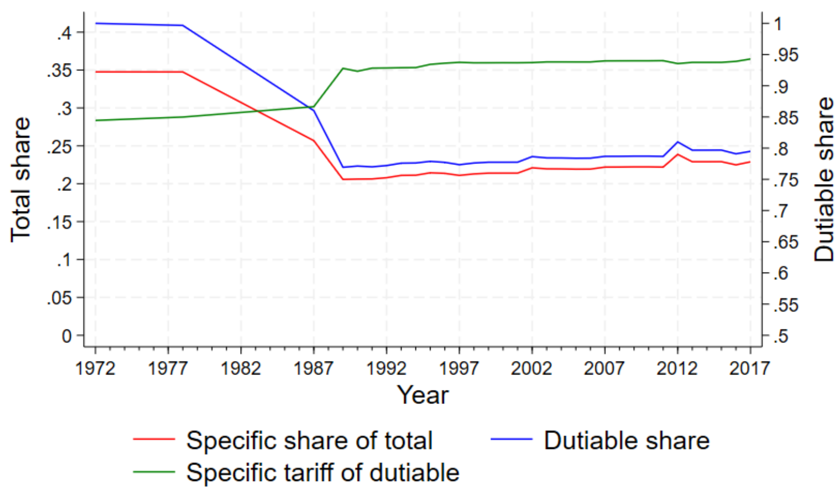
The Uruguay Round column 1 reductions, implemented during the late 1990s and early 2000s, see the share of dutiable column 1 tariff lines fall by roughly one-third, to approxi-

[61] Feenstra et al. (2002) provide tariff-line level unit values, but this does not allow calculation of variety-level AVEs. This is important in the presence of specific tariffs due to the considerable variation in unit values, and hence AVEs, across exporters of a given tariff-line-level good.

Figure G.1: Prevalence of Specific Tariffs over Time: 1972-2017



(a) Column 1



(b) Column 2

Notes: Specific share of total (dutiable) is share of tariff lines (dutiable tariff lines) with specific tariffs, measured on left (right) y-axis. Dutiable share is share of imports with any tariff, measured on right y-axis. To ensure full coverage of column 2 tariff lines, panel (b) uses digitized 1972, 1978, and 1987 tariffs for the TSUSA period.

mately 65%.[62] The share of lines with specific tariffs falls slightly, to 10.5%, but the liberalization is disproportionately tilted towards ad valorem tariff lines, such that the share of dutiable tariff lines with specific tariffs actually increases slightly, to 17%. The post-Uruguay

[62]While the Uruguay Round negotiations were over MFN tariff bindings, the resulting US bindings were sometimes below the MFN applied tariff and led to reductions in US column 1 tariffs.

period is characterized by broad stability in both the share of dutiable tariff lines and the share of tariff lines with specific tariffs.[63]

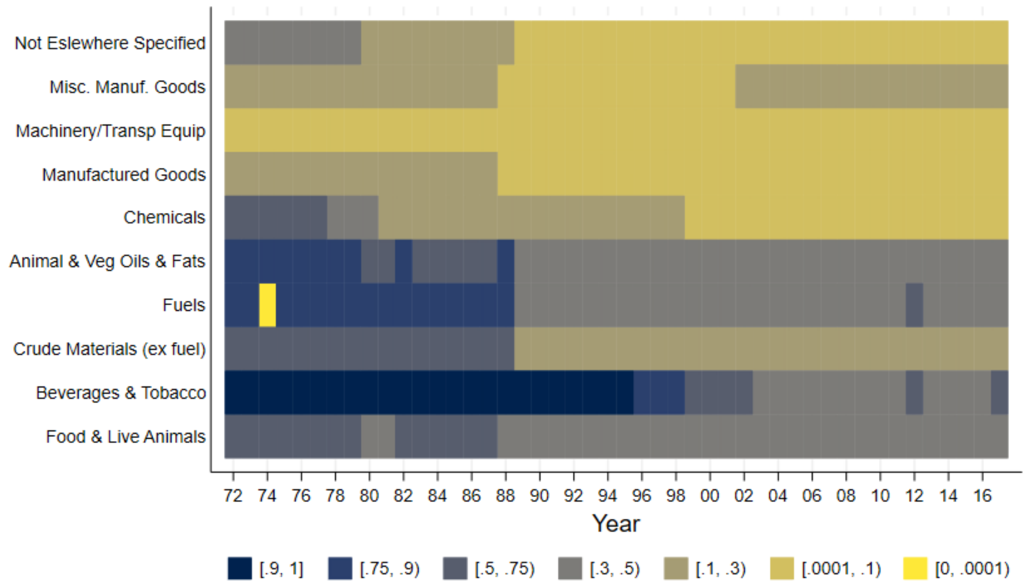
To explore heterogeneity in this process, Figure G.2 depicts the sectoral composition of specific tariffs across one-digit SITC sectors by year. Following Figure G.1, panels (a) and (b) of the figure depict the sectoral composition of column 1 and column 2 specific tariffs, respectively. For both categories, manufacturing tends to report relatively few specific tariffs compared to agriculture and raw materials. The 1972 share of dutiable column 1 tariff lines covered by specific tariffs in the manufacturing sector varies between 5% in "Machinery & Transport Equipment" and 30% in "Manufactured Goods". In contrast, the share for raw materials varies between 50% in "Chemicals" and 90% in "Fuels", while the share in agriculture varies between 58% in "Food & Live Animals" and 100% in "Beverages & Tobacco".

The evolution of specific tariffs over time described in Figure G.1 above is also evident in Figure G.2. In 1980, the first year of Tokyo Round tariff phase-in, the share of column 1 lines with specific tariffs falls by 10 percentage points in "Chemicals" and by approximately 5 percentage points in "Food & Live Animals", "Crude Materials", "Animal & Vegetable Fats & Oils", and "Manufactured Goods". Even more extreme are the changes when transitioning from the TSUSA to HS tariff code in 1989, at which point the share of lines accounted for by specific tariffs falls by 50 percentage points for "Fuels", 27 percentage points for "Crude materials", 33 percentage points for "Animal & Vegetable Fats & Oils", and 9 percentage points for "Food & Live Animals". The Uruguay Round tariff liberalization further reduces the prevalence of specific tariffs, albeit somewhat less dramatically.

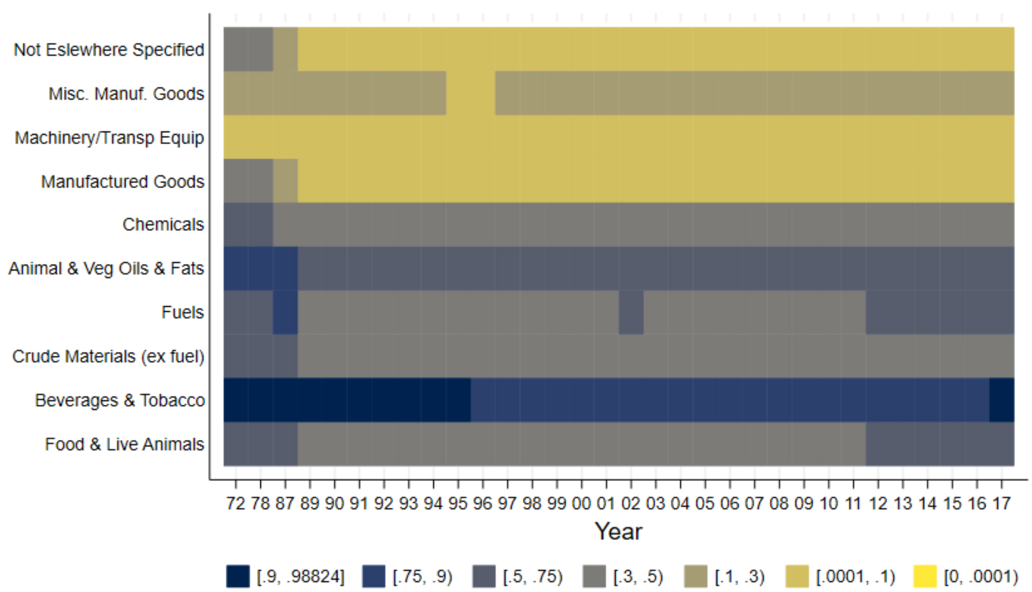
Overall, while the process is both non-linear and varied across sectors, the prevalence of specific tariffs has broadly declined over time as a share of tariff lines. In turn, the way specific tariffs impact the US tariff schedule in the HS era of relatively low and stable inflation is different to the 1970s and 1980s. Figures G.3-G.4 show that specific tariffs and inflation erosion matter little for explaining changes in aggregate or variety-level AVEs over time during the HS era. Instead, fluctuations in the aggregate AVE are closely related to the Uruguay Round tariff cuts as well as formation of US free trade agreements (FTA), especially NAFTA in the mid-late 1990s. These led to falling statutory tariffs on FTA partners during the FTA phase-in period, but FTA partners also began exporting new products to the US and there was a shift in the composition of US imports towards FTA partners.Indeed, only

[63]Figure G.7 at the end of this section modifies Figure G.1 by showing the shares of imports rather than tariff lines. The patterns are broadly similar.

Figure G.2: Sectoral Prevalence of Specific Tariffs over Time: 1972-2017



(a) Column 1



(b) Column 2

Notes: Figure shows share of dutiable tariff lines with specific tariff in one-digit SITC sectors. To ensure full coverage of column 2 tariff lines, panel (b) uses digitized 1972, 1978, and 1987 tariffs for the TSUSA period.

about 40% of the variation in the aggregate AVE stems from variation in the AVE effect and only 60% of the variation in the AVE effect is explained by variation in the statutory ad valorem tariff effect.[64]

While the importance of specific tariffs and inflation erosion in explaining the time-series variation AVEs has fallen notably in the HS era, many applications rely on cross-sectional tariff variation as a source of identification. To this end, the cross-sectional variance of the good-level AVE can be written as

$$\text{var}\left(\tau_{gt}+\frac{f_{gt}}{p^*_{gt}}\right)=\text{var}\left(\tau_{gt}\right)+\text{var}\left(\frac{f_{gt}}{p^*_{gt}}\right)+2\text{cov}\left(\tau_{gt},\frac{f_{gt}}{p^*_{gt}}\right). \tag{G.1}$$

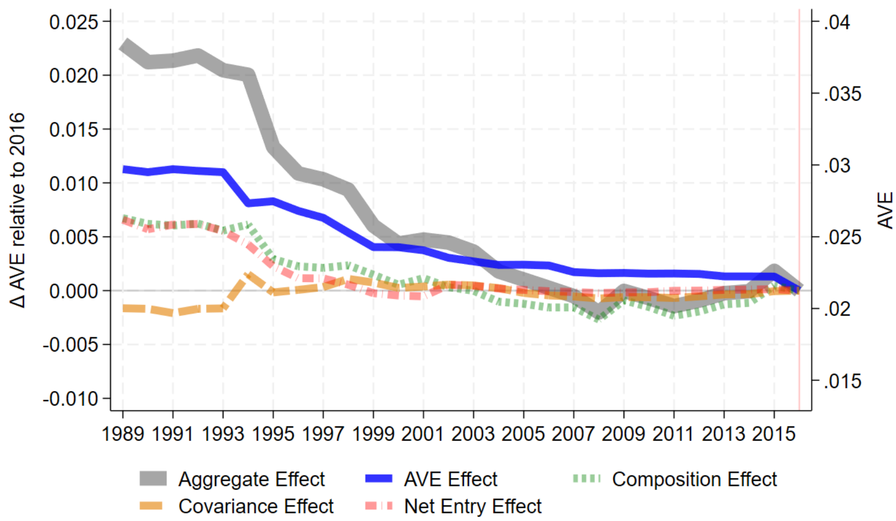
In this context, the relevant measure is the share of cross-sectional variance in tariff-line AVEs attributable to each of the various components: ad valorem tariffs, specific tariff AVEs, and their covariance. Figure G.5 depicts these shares for each year between 1972 and 2017.

Several interesting facts emerge from the figure. First, while the variance of ad valorem tariffs typically accounts for 40% to 60% of the cross-sectional AVE variance, the contribution of specific tariffs is substantial: the share of the variance accounted for by specific tariff AVEs generally lies between 30% and 50%. Second, these shares vary non-monotonically and substantially over time. Indeed, for a period of roughly a decade, the variance of specific tariff AVEs explains a *larger* share of the cross-sectional AVE variance than ad valorem tariffs. This is true when our sample period begins in 1972. But inflationary erosion of specific tariff AVEs during the 1970s, and the conversion of many specific tariffs to ad valorem tariffs in 1980 and 1989 upon transition to the HS code, reduces the importance of specific tariffs and increases the importance of ad valorem tariffs. However, these shifts reverse during the mid-late 1990s and early-mid 2000s when the disproportionate focus of the Uruguay Round liberalization on ad valorem tariffs again affords a larger role for specific tariff AVEs than ad valorem tariffs. Price increases in the late 2000s and early 2010s again reduce the importance of specific tariff AVEs and increase the importance of ad valorem tariffs.

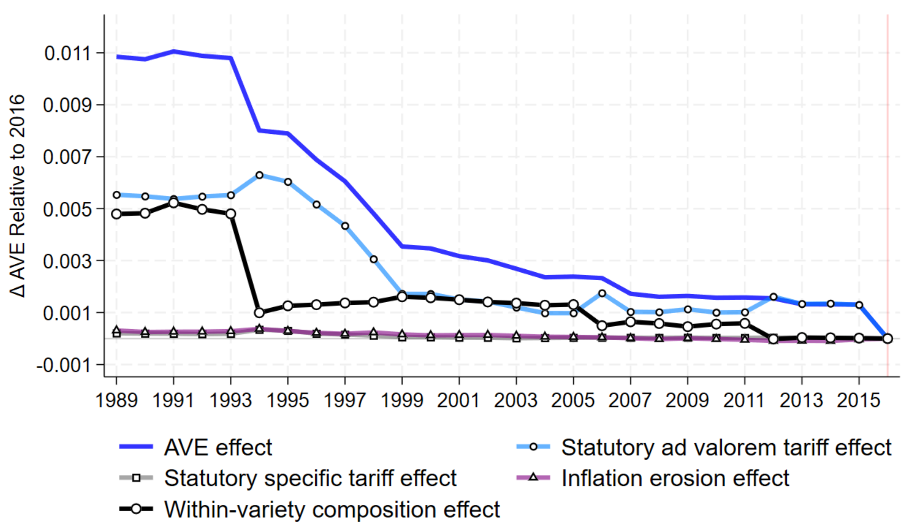
Cross-sectional variation in specific tariff AVEs can stem from differences in statutory specific tariffs or from variation in prices. Figure G.6 shows that the majority of the cross-sectional variation in specific tariff AVEs is driven by the variance of tariff-exclusive prices,

[64] The "within-variety composition effect" is defined as $\Delta AVE_{vgt} - \sum_r s_{vgrt_0}\left[\Delta\tau_{vgrt_0}+\frac{\Delta f_{vgrt_0}}{p^*_{vgrt_0}}-\frac{\Delta p^*_{vgrt}}{p^*_{vgrt_0}}STS_{vgrt}AVE_{vgrt}\right]$ where the share s_{vgrt_0} for rate provision code r is the share of base year t_0 imports of good g from country v.

Figure G.3: AVE Decompositions in the HS Era



(a) Aggregate AVE decomposition

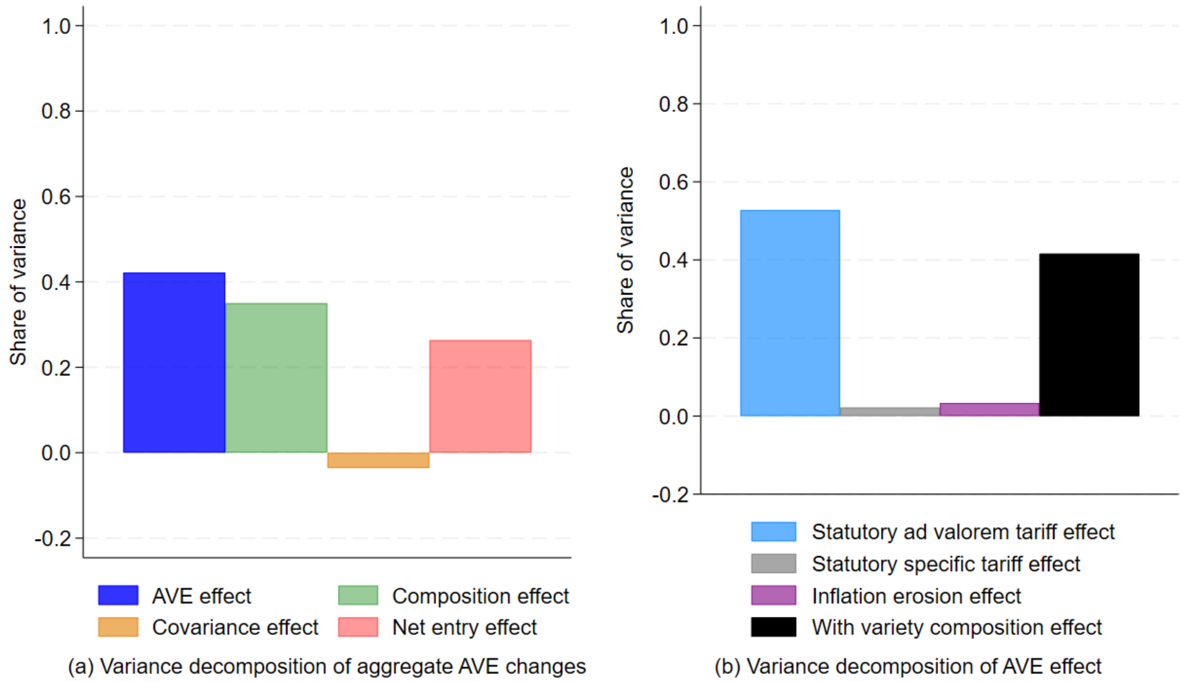


(b) Variety-level decomposition of AVE effect

Notes: Panel (a) shows the decomposition of changes in the aggregate AVE detailed in equation (2). Panel (b) shows the decomposition of the AVE effect in equation (2) and panel (a) by decomposing changes in variety-level AVEs using equation (3) and aggregating using the variety-specific shares in equation (2). This aggregation ensures the share-weighted components from equation (3) sum to the AVE effect in equation (2) and panel (a). Both panels use a base year of $t_0 = 2016$, indicated by a red vertical line.

rather than the tariffs themselves. This further underscores the importance of understanding the sources of variation in AVE, as prices are an endogenous object and not a policy variable.

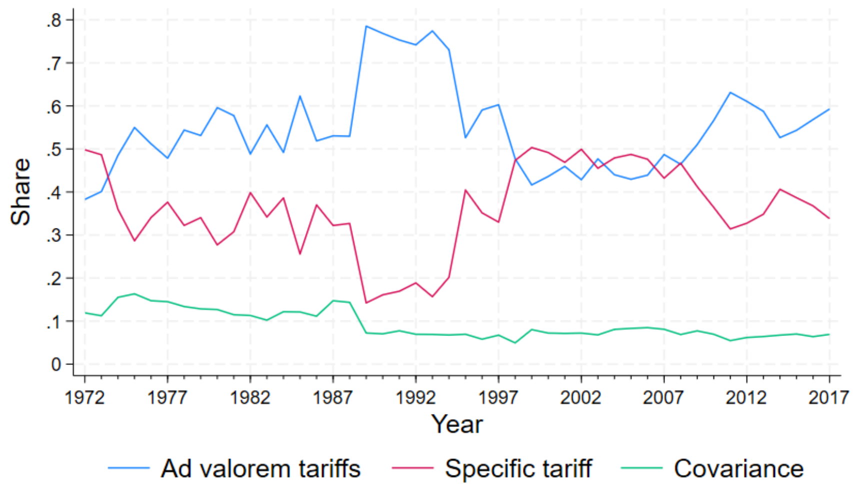
Figure G.4: Sources of Variation in Aggregate AVE Changes in the HS Era



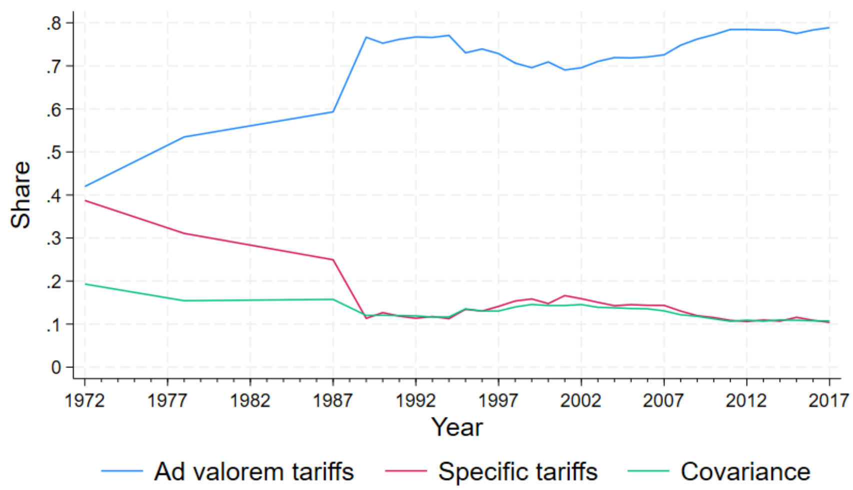
Notes: Figure decomposes the variance of the Aggregate Effect in panel (a) of Figure G.3 and the AVE effect in panel (b) of Figure G.3. It decomposes these into the variance of the constituent components in the respective decompositions shown in Figure G.3, and the pairwise covariance between each of the constituent components. For illustration, the covariance term between two constituent effects is equally allocated between the two effects.

These facts have important implications for empirical analysis relying on cross-sectional AVE variation. While only 15%-20% of dutiable column 1 tariff lines have specific tariffs in the HS period, specific tariff AVEs often explain nearly as much of the cross-sectional variance in the US tariff code as ad valorem tariffs. For analyses exploiting cross-sectional variation in AVE rates in the 21^{st} century, between 35% and 50% of the variation still stems from variation in specific tariff AVEs, with much of this variation reflecting differences in prices rather than differences in tariffs themselves. This is especially relevant for empirical analyses that use initial tariffs to instrument for subsequent tariff changes or, as with free trade agreements, measure tariff exposure as the elimination of tariffs. In such cases, variation in measured liberalization across industries will depend heavily on variation in specific tariff AVEs and, in turn, on variation in tariff-exclusive prices rather than variation in actual tariff policy.

Figure G.5: Cross-sectional Variance Decomposition of US Tariff Code: 1972-2017



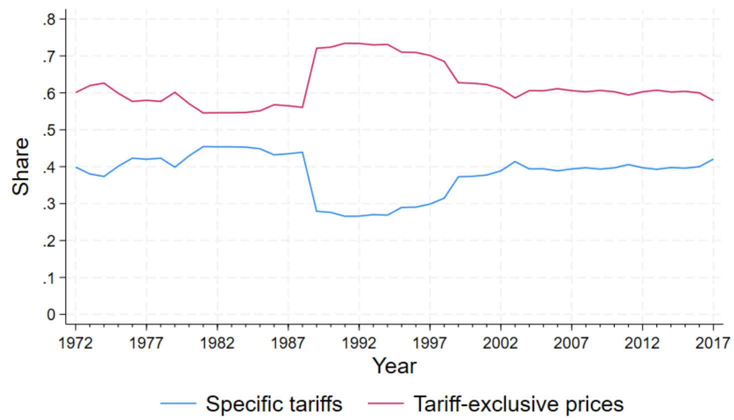
(a) Column 1



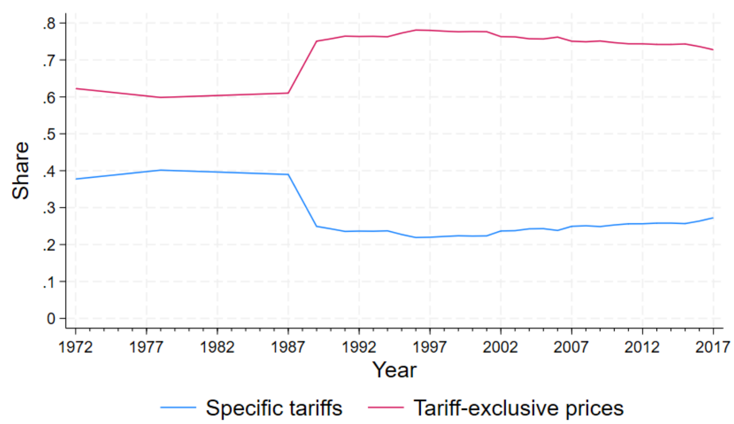
(b) Column 2

Notes: Figure decomposes cross-sectional variance of good-level AVE for dutiable goods in each year using equation (1).

Figure G.6: Cross-sectional Variance Decomposition of Specific Tariff AVE: 1972-2017



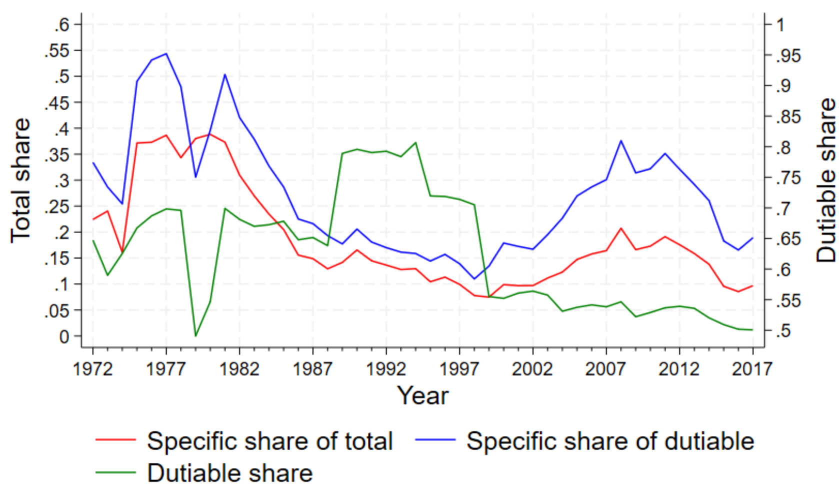
(a) Column 1



(b) Column 2

Notes: Figure decomposes cross-sectional variance of good-level specific tariff AVE in each year for goods with specific tariffs using variant of (G.1): $\text{var} \ln\left(\frac{f_{gt}}{p^*_{gt}}\right) = \text{var} \ln\left(f_{gt}\right) + \text{var} \ln\left(p^*_{gt}\right) - 2\text{cov}\left(f_{gt}, p^*_{gt}\right)$. Given a negative covariance, the shares for specific tariffs and tariff-exclusive prices are normalized to sum to 1.

Figure G.7: Prevalence of Specific Tariffs over Time: 1972-2017



(a) Column 1

Notes: Specific share of total (dutiable) is share of column 1 imports (dutiable imports) with specific tariffs, measured on left (right) y-axis. Dutiable share is share of column 1 imports with any tariff, measured on right y-axis.