Trade and Inequality: Revisiting the 1980s with a New Approach and New Data*

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October 2, 2023

Preliminary and Incomplete

Abstract

Comparatively little work exists exploring the impacts of trade on pre-1990 US labor market outcomes, in stark contrast to post-1990 outcomes. This, in part, reflects a lack of pre-1990 disaggregate tariff data. To address this, we reconstruct annual legislated US tariffs between 1972 and 1988. Our new data reveal a stark two-phase liberalization. Specific tariffs account for one-third of all tariffs at the beginning our sample and facilitate an "accidental liberalization" between 1972 and 1979 as inflation substantially erodes the ad valorem equivalent of specific tariffs. The accidental liberalization is similar in magnitude to the 1980-1988 reduction in legislated tariffs resulting from Tokyo Round GATT negotiations. Based on these historical insights, we develop a novel instrument for tariff reductions in the post-Tokyo era to explore the 1980s relationship between labor market outcomes and tariffs. Our analysis reveals a substantial

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

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role for industry-level tariff cuts in driving income inequality, with more exposed industries experiencing relative increases in skill-intensity. Combined with additional results that industries responded to larger tariff cuts through increased patenting and investment, our analysis points to an important role played by trade-induced skill-biased technological change.

1 Introduction

The past 15 years have witnessed an explosion of influential work on the distributional effects of international trade in the United States. Scholars have documented large and persistent consequences of trade on US workers and firms, with differential exposure driven by geographic location (Autor et al., 2013), industry of employment or production (Pierce and Schott, 2016), or some combination of the two (Hakobyan and McLaren, 2016). In documenting these effects, this literature has overwhelmingly focused on the past 30 years. This is perhaps natural given the rapid growth of trade in the post-1990 era of "hyper-globalization".[1] Indeed, trade liberalization has been a central component of US foreign policy through the Uruguay Round of the General Agreement on Tariffs and Trade (GATT) and the creation of the World Trade Organization (WTO) in 1994, the proliferation of free trade agreements throughout the 1990s and 2000s, and the accession of economic powers China and Russia to the WTO in the early 2000s. However, this focus comes with an important empirical caveat: by 1990, much of the post-war US trade liberalization was already complete. Over the past century, the US ad valorem equivalent (AVE) tariff, defined as the ratio of duties collected to total imports, peaked at approximately 20% following the Tariff Act of 1930. By 1990, this had fallen to 3.3%.[2] Thus, much of the economic impact of trade on the US economy was likely felt in the decades prior to those emphasized by scholars.

One important reason for the relative lack of work on prior decades is a lack of product-level import tariff data, which only become available with the advent of the Harmonized System (HS) in 1989. In this sense, the argument made two decades ago by Anderson and van Wincoop (2004, p.693) that "the grossly incomplete and inaccurate information on policy barriers available to researchers is a scandal and a puzzle" remains true to this day. In this paper, we begin to address this shortcoming by developing an algorithm to reconstruct annual product-level legislated US tariffs between 1972 and 1988, a period when the US aggregate AVE tariff fell by roughly 25% *more* than during the "hyper-globalization" years between 1990 and 2016. Our procedure relies only on publicly available data used extensively by trade economists and yields highly accurate tariff estimates covering the vast majority of US imports. Indeed, we demonstrate that our algorithm yields nearly identical rates to those which would be obtained by a perfect digitization of PDF tariff schedules. These data constitute the first publicly available dataset of annual legislated US tariffs in this period.

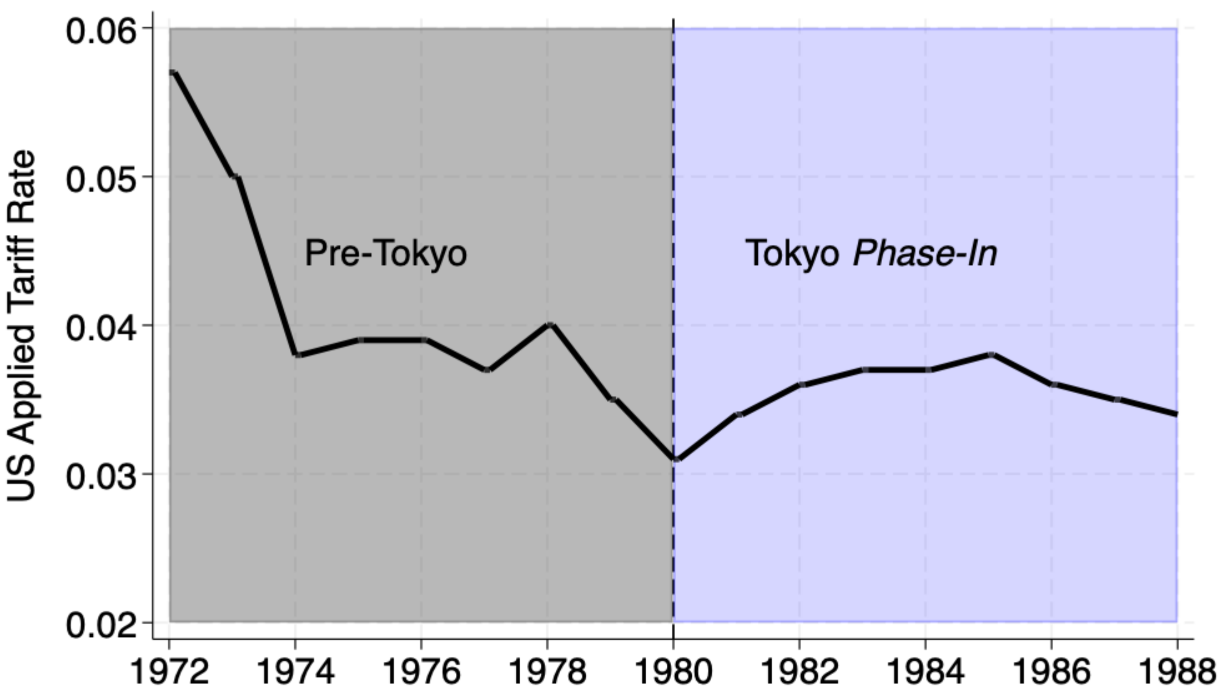
To highlight the importance of these data, we undertake two exercises. First, we use the data to explain a puzzle in the evolution of US tariffs in this era. The Tokyo Round of the GATT was

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

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

completed in 1979, with negotiated tariff reductions 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 1967 Kennedy Round were completed in 1972.[3] One would thus expect relative stability in average US tariffs throughout the 1970s and a subsequent decline throughout the 1980s. However, Figure 1 depicts the US aggregate AVE tariff as reported by the US International Trade Commission (USITC) and shows that the *opposite* is true. That is, the average tariff level fell throughout the 1970s, and despite the successful completion of the Tokyo Round in 1979, average tariffs *rose* slightly in its wake.

Figure 1: Aggregate US Ad Valorem Equivalent Tariff Rate



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

Using our estimates of legislated tariffs, we show that this aggregate pattern is the product of two distinct changes. First, between 1972 and 1979 there is an "accidental liberalization" caused by the inflationary erosion of protection afforded by specific (per-unit) tariffs. As prices rise, the ad valorem equivalent level of protection afforded by nominally defined tariffs declines. Specific tariffs

[3]The establishment of the Generalized System of Preferences (GSP) program in 1974 eliminated tariffs on a wide range of products for many developing economies. While we show below that GSP likely played some role in the pre-Tokyo AVE decline, in particular between 1973 and 1974, it cannot explain the majority of the decline.

account for more than one-third of all tariff lines at the beginning of our sample and ad valorem equivalent rates on goods facing specific tariffs fall from 7% to 3% as prices rise throughout the 1970s. This drives the majority of the aggregate decline shown in the figure, with the remainder accounted for by the introduction of the Generalized System of Preferences program in 1974.[4] Second, the Tokyo Round negotiated tariff reductions are phased in between 1980 and 1988 as a step-wise decline in annual tariff rates. However, this product-level liberalization is masked by the changing composition of imports among continually imported products. Holding import shares fixed among continuing products, we find that tariffs are 27% lower in 1979 than in 1972 and fall another 25% with the tariff phase-ins scheduled under the Tokyo Round.

As a second exercise, we revisit the literature focused on the pre-1990 impacts of trade on inequality. The enormous influence of recent work on the distributional effects of trade noted above was highly improbable from the vantage point of the late 1990s. At that time, the frontier of empirical trade research was quite far from both the theoretical predictions of standard trade models and from popular discontent regarding the passage of high-profile trade agreements such as NAFTA and the Uruguay Round.[5] Indeed, the literature analyzing the effects of trade on labor market outcomes in the era largely agreed that trade played relatively little role in explaining rising equality throughout the 1980s. Instead, it concluded that skill-biased technological change accounted for the bulk of the increase.

One potential cause of the gap between data and both classical theory and public perception was proposed by Slaughter (2000). He concluded that despite years of research by leading trade economists, "more work needs to link the various exogenous forces attributable to international trade to actual product-price changes ... the literature to date has made substantial progress understanding how to relate a given change in relative product prices to changes in relative factor prices ... it has made less progress understanding whether these product-price changes have any thing to do with international trade". In the terminology of modern empirical microeconomics, the literature struggled with the fundamental issue of identification. Work exploring the relationship between trade and labor market outcomes prior to 1990 was largely conducted either as a test of the factor proportions theory, with less emphasis placed on causal effects more generally (Lawrence and Slaughter, 1993; Berman et al., 1994; Borjas et al., 1997) or as a correlational analysis between trade flows and wages or employment (Kletzer, 2000; Lovely and Richardson, 2000). Very little work exists attempting to establish a causal relationship between tariff changes and labor market outcomes in this era. The causal effects on the US economy of the 1979 Tokyo Round of the GATT,

[4] The importance of specific tariffs for tariff levels in this era was noted in more aggregate data by Van Cott and Wipf (1983).

[5] Such unease regarding the benefits of trade was famously captured by Presidential candidate Ross Perot's prediction of a "giant sucking sound" following NAFTA as jobs left the US for Mexico, or the "Battle of Seattle" protests at the 1999 WTO Ministerial Conference.

in particular, have received virtually no attention from scholars.[6]

We leverage our data to address this issues directly. In particular, we exploit the fact that the Tokyo Round tariff reductions largely proceeded according to the pre-determined "Swiss Formula" to construct a novel instrument for tariffs that avoids standard endogeneity concerns associated with relating tariff changes to labor market outcomes. The instrument strongly predicts import values following the Tokyo Round, but is only weakly correlated with export values, suggesting that it is not merely a proxy for increased trade more broadly. Using this instrument, we show that industries more exposed to tariff reductions between 1979 and 1988 witnessed an increase in the value of shipments but a reduction in share of that value accounted for by payments to labor. Further, we find a relative reduction in the share of the wage bill accounted for by production workers, driven by a relative increase in employment of non-production workers. We demonstrate that an approach failing to account for the endogeneity of tariff reductions does not find a statistically significant relationship between tariffs and labor market outcomes in the 1980s, or even between tariffs and import values. Thus, a final contribution of the paper is an improved understanding of the relationship between trade liberalization and labor market outcomes in the 1980s based on a novel identification strategy relying on the details of the Tokyo Round tariff code. We believe this approach will be useful in exploring the economic response to trade in this era more broadly.

Our work is related to four distinct strands of literature. First, as noted above, we contribute to the large literature focused on the relationship between trade and rising inequality throughout the 1980s. The majority of this literature took the Heckscher-Ohlin model as its starting point and explored the relationship between changes in goods prices and changes in factor prices as a test of the Stolper Samuelson theorem. Much of this took the form of "price regressions", in which changes in industry product prices are regressed against industry factor cost shares, with the point estimates yielding the changes in factor prices implied by the observed changes in product prices under the assumption of zero profits (Leamer, 1998; Feenstra and Hanson, 1999). This literature implicitly assumed perfect factor mobility between sectors and thus focused on changes in a single, economy-wide wage for skilled and unskilled labor. Our approach, conversely, focuses on relative changes between industries. While this difference-in-differences approach prevents us from making statements about aggregate inequality, it does not require the assumption of factor mobility or perfect competition embedded in factor proportions models. Further, the magnitude of relative changes in skill pay differentials across industries that we document is large – an interquartile reduction in AVE tariff levels corresponds to an increase in relative to payments to non-production workers of 0.035 log points relative to a 1979-1987 increase of 0.2.

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

Second, we add to the numerous studies focused on quantifying the effects of trade policy on US domestic economic outcomes more generally. Studies of the 1989 Canada-US Free Trade Agreement (Trefler, 1993; Kovak and Morrow, 2022), NAFTA (Caliendo and Parro, 2015; Hakobyan and McLaren, 2016), the US granting PNTR to China in 2001 (Pierce and Schott, 2016; Handley and Limao, 2017; Greenland et al., 2019), the 2001 Bush steel tariffs (Cox, 2022; Lake and Liu, 2021), and the Trump-era tariffs (Fajgelbaum et al., 2020; Flaaen and Pierce, 2021) all rely on discrete changes in disaggregate tariff data to establish causal links between trade policy and economic outcomes. In addition to exploring an earlier era than these papers, our paper is distinct in providing an instrument for a multilateral liberalization that avoids standard endogeneity concerns. Further, in hopes of facilitating subsequent research in this era, we also construct the first publicly available dataset of TSUSA legislated tariffs at the annual level, designed to complement the US import data constructed by Feenstra (1996).[7] Given the considerable growth in imports, substantial declines in tariffs, and labor market turmoil during this time period, we believe these data and techniques will prove valuable for other researchers.

While not focused on a specific policy shift, the paper most similar to ours in its emphasis on trade and labor market outcomes in this era is Batistich and Bond (2023), who examine the effect of rising imports from Japan on local labor market outcomes between 1970 and 1990. Using Japanese exports to third markets as an instrument for exports the US in the spirit of Autor et al. (2013), they find that rising import growth from Japan corresponds to relative reductions in Black manufacturing employment and labor force participation and declining relative wages for Black males. We differ from them in exploiting the entirety of the tariff code, rather than focusing on Japan specifically. Further, our identification relies on the exogeneity of pre-existing tariff rates and the parameters of the "Swiss Formula", rather than any particular assumption about export growth in foreign markets.

Third, our study contributes to the literature focused on the impacts of additive trade barriers. Hummels and Skiba (2004) and Feenstra and Romalis (2014) emphasize the importance of per-shipment trade costs in determining both trade flows and export quality. Irarrazabal et al. (2015) detail a procedure to estimate such costs, while Sørensen (2014) shows that the gains from trade are larger in the presence of these types of trade barriers. Still other studies emphasize the importance of per-unit trade costs in shaping tariff levels and import patterns during the first half of the 20$^{\text{th}}$ century US (Crucini, 1994; Irwin, 1998; Bond et al., 2013; Greenland and Lopresti, 2022). In this paper, we show that such trade costs are quantitatively important for trade flows far more recently than has been previously demonstrated. Specific tariffs cover more than one-third of products in 1972, and more than 25% of goods on the eve of the transition to the HS system.[8]

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

[8] As of 2020, specific tariffs are still a highly important form of protection in US agriculture. Moreover,

Finally, our paper contributes to the body of work aimed at expanding and improving access to historical tariff data. In ongoing work, Acosta and Cox (2022) digitize US tariff schedules following major trade agreements, beginning with the Smoot-Hawley Act of 1930 to document and explore persistent regressivity, with higher-end goods facing lower tariff levels. Their sample covers a wider range of GATT rounds than ours, but is not constructed at the annual level. To the extent that legislated rates may have changed between the Kennedy and Tokyo Rounds – for instance, if applied tariffs were adjusted below negotiated levels (Nicita et al., 2018) – our application requires annual data. Further, in addition to producing annual data on legislated tariffs for the 1972-1988 period, the algorithm we construct and detail in this paper can be applied in other settings. The only requirement for estimating per-unit and ad valorem tariffs is product-level import values, duties, and quantities at the same level that tariffs our set. Such data are widely available.

The paper proceeds as follows. Section 2 describes the data and algorithm used to reconstruct the TSUSA tariffs from 1972-1988 and evaluates the accuracy of the estimates. Section 3 provides an explanation for the puzzle surrounding the disconnect between the aggregate US AVE tariff rate and conventional wisdom about US tariff policy between the Kennedy and Tokyo Rounds. Section 4 explores the relationship between tariff levels and industry-level outcomes between 1979 and 1987 using an instrument based on the "Swiss Formula" for tariff liberalization. Section 5 concludes.

2 Data

We begin by discussing construction of a database of annual US legislated tariffs. Our sample begins in 1972, the year in which the US completed phasing in its tariff commitments under the Kennedy Round of the GATT, and ends in 1988, the final year before the US changed tariff classification systems from the Tariff Schedule of the United States Annotated (TSUSA) to the Harmonized Tariff Schedule of the United States (HTS). Currently, no dataset contains annual product-level US legislated tariffs for these years. As a result, prior work using tariffs in this period has generally focused on the AVE tariff rate, dividing duties collected by import values. This can be done using US import data recorded annually in the US Census Annual Import Data Bank (IDB), as extensively documented by Feenstra (1996). However, such an approach ignores the distinction between ad valorem tariffs – specified as a percentage of import value – and specific tariffs – specified per unit of quantity. Researchers pursuing such an approach are thus unable to determine whether changes in AVE rates are driven by changes to legislated tariffs or by endogenous price movements in the presence of specific tariffs. Given the dramatic increases in price levels throughout the era and

some countries remain dependent on specific tariffs for protection across the entire tariff schedule. Over 85% of Switzerland's tariff code, for example, consists of specific tariffs.

the pervasive use of specific tariffs in the US at the time, this is a non-trivial concern.[9] Our data set, which includes annual product-level legislated tariffs – both ad valorem and specific – enables researchers to separately identify these two forces.

2.1 Estimating Tariffs 1972-1988

Rather than manually digitizing the published tariff schedule for each year in our sample, which is a time-intensive and error-prone task, we estimate legislated rates using publicly available data on imports.[10] Specifically, we rely on data from the IDB that we access from The Center for International Data at UC Davis.[11] These data include total duties collected as well as import values, dutiable values, and up to two measures of quantity. The data are highly disaggregate: each observation represents exports by country of a seven-digit TSUSA product to a US customs district in year t subject to tariffs defined by "rate provision code" r.[12]

A rate provision code defines the applicable tariff along two dimensions. The first dimension indicates whether the tariff is duty free, subject to ad valorem tariffs, subject to specific tariffs, or subject to a combination of the two types, known as "compound" tariffs. For non-duty free imports, the second dimension indicates whether the applicable tariff is from column 1 of the TSUSA – the "normal trade relations" or "NTR" tariff – or from column 2 of the TSUSA – the "non-NTR" tariff. For duty-free imports, the second dimension indicates whether the zero duty applies through a form of preferential access, such as the Generalized System of Preferences program. For each year in our sample, we estimate the legislated ad valorem and specific tariffs for each observed rate provision code.

For each rate provision code r in year t, one can express total duties collected on imports of variety v – defined as an exporter-by-seven digit TSUSA product-by-customs district – as

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

Here, DV_{vgrt} and Q_{vgrt} represent, respectively, the dutiable value and quantity of imports of variety v within five-digit good g, the level at which tariffs are defined. Parameters τ_{grt} and f_{grt}

[9]For instance, the US CPI nearly tripled between 1972 and 1988: https://fred.stlouisfed.org/series/CPIAUCSL#0.

[10]Each TSUSA schedule contains roughly 700 pages of tariffs covering up to 7000 products annually.

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

[12]A customs district is actually a pair defined by a district of entry (i.e. the customs district where imports clear US customs) and a district of unlading (i.e. the customs district where the imports are unloaded from the importing vehicle) which may or may not be the same.

represent, respectively, the unobserved tariff line-level legislated ad valorem and specific tariffs. Absent measurement error, one could simply calculate these legislated tariffs using the IDB import data. Specifically, using data on total duties, quantity, and dutiable value for any variety-level observation, one would recover ad valorem tariffs τ_{grt} and specific tariffs f_{grt} as

$$\tau_{grt} = \frac{TD_{vgrt}}{DV_{vgrt}} \tag{2}$$

$$f_{grt} = \frac{TD_{vgrt}}{Q_{vgrt}}. \tag{3}$$

Similarly, one would recover the ad valorem and specific components of compound tariffs by solving the two simultaneous equations defined by any two non-collinear variety-level observations:[13]

$$TD_{vgrt} = \tau_{grt} DV_{vgrt} + f_{grt} Q_{vgrt}. \tag{4}$$

However, multiple forms of measurement error make direct calculation infeasible. The first is that multiple years of IDB data do not report the unit of quantity. This is problematic for estimating specific tariffs because the IDB data often report quantity values for two units of quantity, with each unit generating its own estimated specific tariff via (3) or, in the case of compound tariffs, its own estimated specific and ad valorem tariffs via (4).[14] In such cases, one would need to solve equation (3) or (4) separately for each unit of quantity. Only one of the two units would yield the same tariffs across the different variety-level observations and these would be the correct tariffs.

Additional sources of measurement error affect all goods regardless of tariff type. First, countries may move between rate provision types within a year and hence face different tariffs during the year.[15] This introduces estimation error, as the IDB records imports annually. Second, the IDB data round import values and quantities to the nearest integer. Third, database entry error may occur – occasionally, for instance, the units of quantity may be recorded inconsistently within a rate code.

Thus, rather than *calculating* the legislated rates, we can only *estimate* the five-digit tariffs τ_{grt} and f_{grt} with error. In particular, we employ a three-step iterative process to produce our final sample. The first step provides initial estimates of the tariffs. The second step refines these

[13] In 0.2% of observations covering 0.1% of imports by value we do not have two non-collinear variety-level observations *vgrt* within the set of good-by-rate provision code-by-year observations *grt*. Thus, we cannot estimate tariffs for these observations.

[14] For example, the five-digit TSUSA good 37315 covers non-ornamented, wool neckties for men and boys. In 1978, the published TSUSA schedule indicates that the compound tariff was 10.5% plus $0.375 per pound. However, the 1978 IDB data report quantity values for two different units of quantity without specifying the unit. It is thus not possible to determine directly which quantity value corresponds to pounds.

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

estimates by removing outlier observations plausibly influenced by the measurement error issues described above. The final step jointly re-estimates the tariffs with data pooled across multiple years. For this step, we estimate the tariffs of five-digit goods for "spells" of consecutively observed years in which the variation in the good's tariff is sufficiently small. Without smoothing annual estimates in this way, we would mistakenly interpret minor year-to-year fluctuations in estimated legislated tariffs as actual legislated tariff changes. We now turn to each step in detail.

Step 1: Initial Estimates Initial estimates of tariffs τ_{grt} and f_{grt} come from estimating the following equation separately for each product g, rate provision code r, and year t over the period 1974-1988:

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

For goods facing only specific tariffs, we impose $\tau_{grt} = 0$ when estimating the equation, while for ad valorem-only goods, we set $f_{grt} = 0$. As noted above, we estimate equation (12) separately for each unit of quantity when the IDB data provide two units of quantity. Assuming each unit yields non-negative estimated tariffs $\hat{f}_{grt}$ and $\hat{\tau}_{grt} > 0$, we use the estimates that yield the higher R^2 as the initial estimates.[16]

Note that the IDB data do not contain total duties for 1972 and 1973. As a result, we cannot use our algorithm to estimate the 1972 and 1973 tariffs. However, as we will show below, our results confirm that legislated tariffs do not change between the Kennedy and Tokyo GATT rounds of tariff negotiations. Thus we are able to backfill the 1972 and 1973 tariffs with our 1974 tariff estimates.

Step 2: Initial Estimate Refinement The first step yields initial estimates $\hat{f}_{grt}$ and $\hat{\tau}_{grt}$ for each five-digit good g. As a test of the accuracy of these estimates, we compare ad valorem equivalent tariffs implied by the IDB data, AVE^{Data}_{vgrt}, to our estimates of these, $\widehat{AVE}_{vgrt}$:

$$AVE^{Data}_{vgrt} \equiv \frac{TD_{vgrt}}{DV_{vgrt}}$$
$$\widehat{AVE}_{vgrt} \equiv \hat{\tau}_{grt} + \frac{\hat{f}_{grt}}{p^*_{vgrt}}$$

[16] In some infrequent cases, goods receive numerous specific duties on components of the individual good. Because the IDB measures at most two distinct quantities, neither digitization nor estimation can recover the true tariff on such goods. Due to the infrequency of this type of good and a concern regarding over-fitting, we do not estimate equation (12) with two units of quantity simultaneously.

where $p^*_{vgrt} = \frac{V_{vgrt}}{Q_{vgrt}}$ is the unit value on the total value of imports V_{vgrt}.[17] We define estimation error as

$$\varepsilon^{Est}_{vgrt} \equiv \ln\left(1 + AVE_{vgrt}\right) - \ln\left(1 + \widehat{AVE}_{vgrt}\right) \tag{6}$$

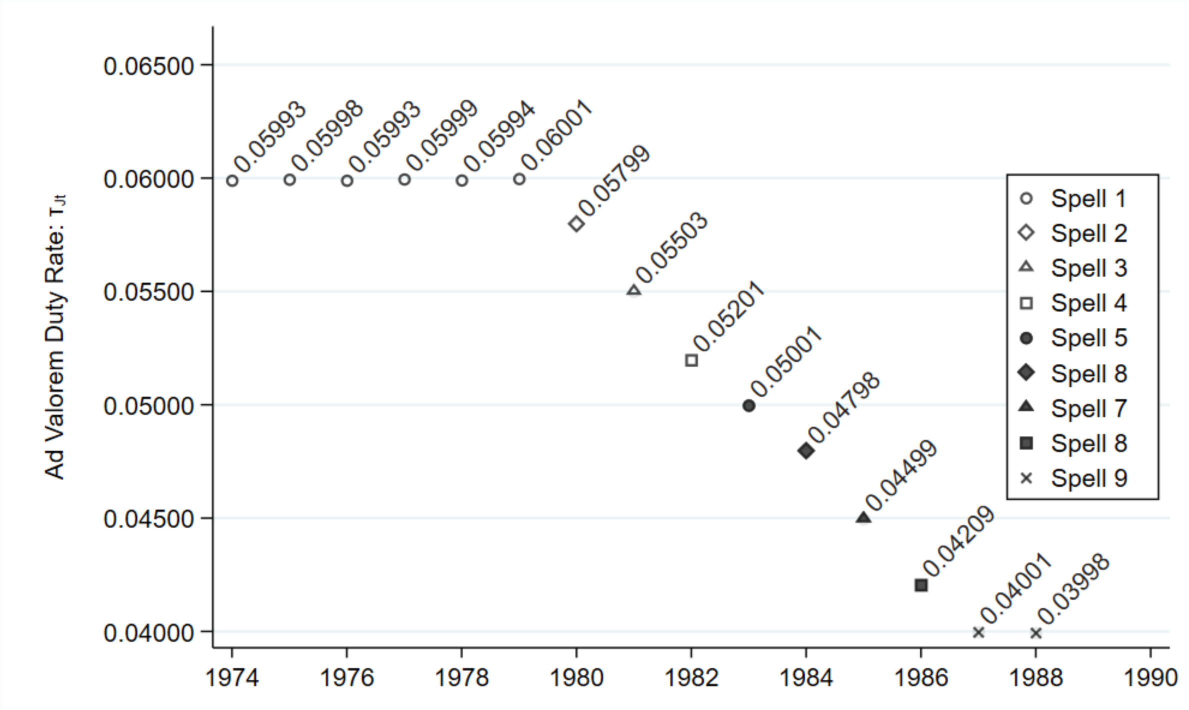
as our measure of estimation error. For good-by-rate provision-by-year observations grt in which the mean absolute estimation error $|\varepsilon^{Est}_{vgrt}|$ exceeds 0.001, we drop the variety v with the largest estimation error from our sample.

Proceeding from the resulting reduced sample, we iterate on the above two steps – estimating tariffs and dropping variety-level observations that drive significant measurement error – up to 10 times. At this point, if we continue to observe a mean absolute estimation error $|\varepsilon^{Est}_{vgrt}|$ that exceeds 0.001, we infer that the error may be driven by individual exporters. We thus drop all observations for an exporter within good g in a given year if the mean absolute estimation error across variety-by-rate provision observations within the exporter-year exceeds 0.001. We then iterate on this step – estimating tariffs and removing any exporter-year observations driving large measurement error – an additional 10 times.

Step 3: Smoothing Annual Estimates Any estimating procedure using reported import data will produce tariff estimates that vary slightly over time, even in the absence of legislated changes. For example, Figure 2 shows that the estimated ad valorem tariff for TSUSA good 11137, salted or pickled herring, takes slightly different values between 1974 and 1979, but that the true legislated value was likely constant at 0.06 in this period before being reduced gradually to 0.04 following the Tokyo Round.

[17] Total imports V_{vgrt} and dutiable value DV_{vgrt} are equivalent for any non-zero τ_{vgrt} or f_{vgrt}. Such cases are the only instances that require p^*_{vgrt}.

Figure 2: Identifying Tariff Spells from Annual Estimates



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

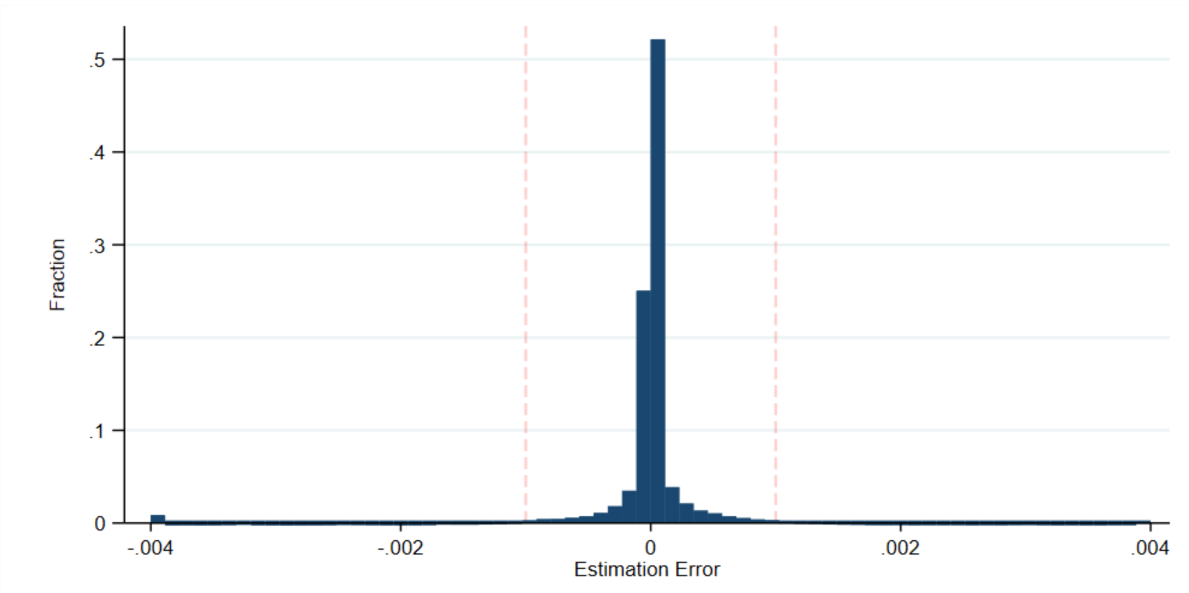
To deal with this issue, we identify "spells" in the data in which the estimated legislated tariffs for a particular good-by-rate provision code suggest that the legislated tariff is constant. Specifically, we define a spell at the good-by-rate provision level gr as a time period between years t_0 and t_1 in which the estimated ad valorem and specific tariffs $\widehat{\tau}_{vgrt}$ and $\hat{f}_{vgrt}$ change by less than 0.015 log points between each year. The two spells in Figure 2 are thus 1974-1979 and 1987-1988. For each spell at the good-by-rate provision code level, we pool all observations across years and repeat steps 1 and 2 above. This yields our final estimates $\widehat{\tau}_{vgrt}$ and $\hat{f}_{vgrt}$ of legislated tariffs τ_{vgrt} and f_{vgrt}.

2.2 Accuracy and Coverage of Estimated Tariffs

Two potential issues arise from estimating legislated tariffs, one related to the quantity of estimates, and one related to their quality. First, our set of estimates may omit a large number of TSUSA products due to the fact that our algorithm drops observations driving large measurement error. Second, our algorithm may produce unreliable estimates. Fortunately, our algorithm performs well along both the quantity and quality dimensions.

In terms of coverage, our estimation algorithm produces a database containing legislated tariffs for 7366 five-digit TSUSA goods. This represents nearly 97% of import value observed in the raw data. To examine the extent of estimation error in our database, we take two approaches. First, we calculate the estimation error ε_{vgrt}, as defined in equation (6). Figure 3 illustrates the distribution of this error. The dashed red lines indicate thresholds of $|\varepsilon_{vgrt}| = 0.001$. Fewer than 3.8% of observations less than and 0.01% of import value lies outside of these thresholds. That is, ad valorem equivalent rates implied by our estimated tariffs very closely match those in the data.

Figure 3: Distribution of Estimation Error



Notes: Estimation error defined in equation (6) at exporter-port-TSUSA 7 digit-rate provision code-year level as log difference between tariff inclusive price and our estimated tariff inclusive price using our tariff estimates from Section 2.1. See main text for further details.

To more directly assess the accuracy of our estimation algorithm, we digitize the fifth supplement of the 1978 TSUSA schedule – the last version published prior to the initial Tokyo Round phase-ins – and compare these digitized tariffs with our estimates. Collapsing over varieties within good-by-rate provision codes yields a sample of 9422 observations, including both column 1 and column 2 tariffs. Based on this sample, and using 1978 unit values, Table 1 reports the import-weighted pairwise correlation between three measures of AVE tariff levels: those in the IDB data, those calculated using our estimates, and those using the digitized tariff schedule. Strikingly, Table 1 shows a near-perfect correlation between our estimated AVE rates, those in the IDB data, and those implied by digitizing the tariff schedule.[18] That is, our algorithm yields estimates that are

[18]This figure omits 83 observations which exhibit nontrivial discrepancies between data and estimation

virtually identical to those that would be obtained by a complete digitization of the tariff schedule.

Table 1: Sample Correlation of Estimated vs Digitized Tariffs

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

Notes: Import weighted pairwise correlation matrix of 1978 good-by-rate level tariffs from the data, estimated via the algorithm detailed above, and digitized.

3 US Trade Policy: 1972-1988

3.1 Decomposing changes in the US aggregate AVE tariff

With our data set constructed, we now turn to an analysis of US trade policy between 1972 and 1988. Figure 1 above highlights two surprising features of the aggregate AVE tariff rate in this period. First, AVE levels declined sharply in the pre-Tokyo Round period despite the fact that NTR rates were unchanged during this period. Second, the aggregate AVE tariff was relatively stable, and in fact *increased* slightly, in the post-Tokyo period despite the negotiated reduction in tariff levels.

To explain these puzzling dynamics, we first note that any change in the aggregate AVE tariff is a product of three different channels: (i) changes in legislated tariff rates, (ii) changes in the AVE of specific tariffs due to price movements, and (iii) changes in the composition of imports along the intensive or extensive margins. To formalize this, we first write the aggregate AVE tariff as the share-weighted average of variety-level AVE tariffs:

$$\begin{aligned} AVE_t =& \frac{\sum_g \sum_v TD_{vgt}}{\sum_g \sum_v M_{vgt}} \\ =& \sum_g \sum_v s_{vgt} \left(\tau_{vgt} + \frac{f_{vgt}}{p^*_{vgt}} \right) \\ =& \sum_g \sum_v s_{vgt} AVE_{vgt} \end{aligned} \tag{7}$$

related to units of quantity in the underlying IDB data.

where s_{vgt} is the year-t variety-level share of imports within good g:

$$s_{vgt} \equiv \frac{M_{vgt}}{\sum_v M_{vgt}}, \tag{8}$$

We then define three mutually exclusive sets of imported varieties: continuing varieties V^C imported in both year t_0 and t_1, exiting varieties V^X imported in t_0 but not t_1, and entering varieties V^N imported in t_1 but not t_0. The change in the aggregate AVE tariff can then be written as:

$$\begin{aligned} \Delta AVE_{t_1} \equiv & AVE_{t_1} - AVE_{t_0} \\ = & \underbrace{\sum_{v \in V^C} s_{vt_0} \Delta AVE_{vgt}}_{\substack{\Delta\text{AVE tariffs,} \\ \text{Import composition fixed}}} + \underbrace{\sum_{v \in V^C} \Delta s_{vgt} AVE_{vgt_0}}_{\substack{\Delta\text{ Import composition (intensive),} \\ \text{AVE tariffs fixed}}} + \underbrace{\sum_{v \in V^C} \Delta s_{vgt} \Delta AVE_{vgt}}_{\substack{\Delta\text{ AVE tariffs,} \\ \Delta\text{ Import composition (intensive)}}} \\ & + \underbrace{\sum_{v \in V^N} s_{vgt_1} AVE_{vgt_1} - \sum_{v \in V^X} s_{vgt_0} AVE_{vgt_0}}_{\Delta\text{ Import composition (extensive)}} . \end{aligned} \tag{9}$$

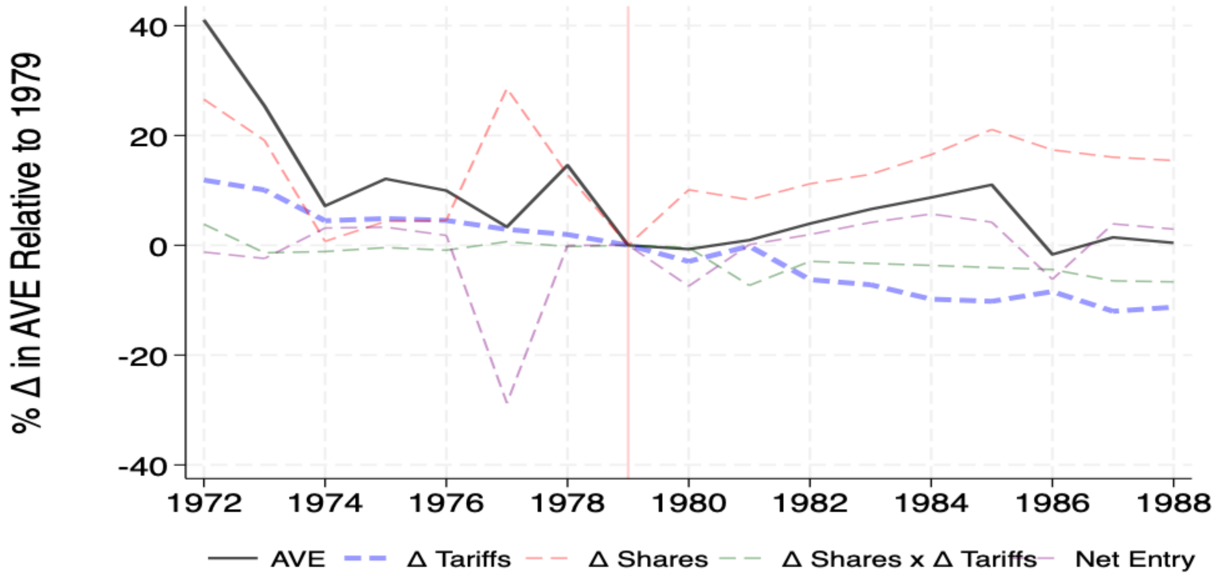
Equation (9) shows that changes in the aggregate AVE tariff consist of changes in variety-level AVE tariffs holding the composition of imports fixed (the first term) and changes in the composition of imports along both the intensive and extensive margins (the last three terms).

Figure 4 illustrates the four components of this decomposition relative to a base year of 1979.[19] The solid black line depicts aggregate AVE, while the dashed lines depict the separate components of the decomposition. Two stark points emerge from the figure. First, the dashed blue line, which depicts changes in variety-level AVE tariffs holding the composition of imports fixed, tracks the aggregate AVE very closely in the pre-Tokyo period. That is, the pre-Tokyo reduction in the AVE tariff is driven almost entirely by variety-level AVE tariff changes rather than by compositional effects.

In contrast, these two series diverge sharply in the post-Tokyo period, implying a role for compositional effects. In particular, consider the dashed red line, which displays changes in the composition of imports across varieties along the intensive margin, holding both legislated tariffs and the variety set constant. Comparing this line to the blue dashed line reveals that increases in AVE tariff through intensive margin compositional effects completely offset the decrease in the AVE

[19]This black line differs slightly from the data presented in Figure 1. While the line in Figure 1 are taken directly from the USITC, the decomposition figure are based only on the tariff data available in our sample. Because our 1972-1973 data are based on the 1974 estimates these data necessarily omit varieties that exit the sample prior to 1974.

Figure 4: Decomposition of Percent Change in Aggregate US AVE Tariff Rate



Notes: Plot based on results from equation (9) where $t_0 = 1979$. Indexed AVE plots ΔAVE_t, other lines plot respective components of equation (9). See main text for further details.

tariff through legislated tariff cuts. That is, while variety-level tariffs do indeed decline, imports shift towards products and exporters with higher initial tariffs, masking the liberalization at the aggregate level.[20]

3.2 Decomposing changes in variety-level AVE tariffs

Figure 4 highlights the role of falling variety-level AVE tariffs in explaining US trade policy in both the pre- and post-Tokyo periods. However, it does not address how this change took place; in particular, it does not explain whether the decline was driven by falling legislated tariffs or rising prices in the presence of specific tariffs. In pursuit of an answer to this question, we decompose variety-level AVE tariff changes into these two components:

[20]The dashed green line provides corroborating evidence for this interpretation: the post-Tokyo composition of US imports shifts towards gods with falling AVE tariffs. As discussed below, the Swiss formula used to determine legislated tariff cuts post-Tokyo implies that goods with high pre-Tokyo tariffs see the largest post-Tokyo cuts.

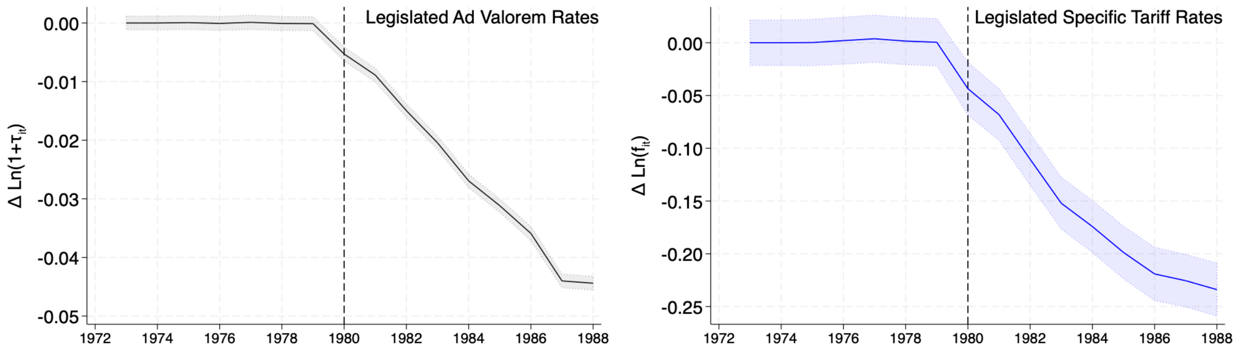
$$\Delta AVE_{vgrt} = \Delta\tau_{grt} + \frac{\Delta f_{grt-1}}{p^*_{vgrt-1}} - \frac{f_{grt-1}}{p^*_{vgrt-1}}\frac{\Delta p_{vgrt}}{p^*_{vgrt}} \tag{10}$$

The first two terms on the right hand side of equation (10) correspond to changes in ad valorem and specific tariffs, respectively. The third term captures changes in AVE through movements in tariff-exclusive prices p^*_{vgrt}. A variety's exposure to price-induced AVE tariff is captured by the size of the specific tariff as a share of its tariff-exclusive price, $\frac{f_{vgrt-1}}{p_{vgrt-1}}$. Intuitively, price increases erode a larger portion of a variety's tariff protection when its specific tariff accounts for a larger share of its tariff-exclusive price. In contrast to the well-documented liberalization of US legislated rates following the Tokyo Round, there is no documented large-scale change in legislated tariffs between the Kennedy and Tokyo rounds. In terms of equation 10, we thus expect the first two terms on the right hand side to play little to no role in the pre-1979 decline in the aggregate AVE tariff.

To confirm this, we use our estimates of legislated tariffs for all five-digit TSUSA products with non-zero tariffs and regress annual product-level log legislated tariffs $\ln(\tau_{grt})$ and $\ln(f_{grt})$ on product fixed effects and year dummies. The point estimates for year fixed effects will thus capture any temporal variation in average legislated variety-level rates. Figure 5 plots the results, with the left and rights panels respectively corresponding to the regression results for ad valorem and specific tariffs, with 1972 as the omitted year. Consistent with our expectations, legislated tariffs are indeed constant in the pre-Tokyo period. This implies that the decline in variety-level tariffs, and thus in the aggregate AVE tariff, in the pre-Tokyo period is driven exclusively by price growth on products with specific tariffs, rather than changes in the underlying legislated tariffs. Conversely, and consistent with Figure 4, legislated tariffs fell substantially in the post-Tokyo period. Relative to the 1972 baseline, ad valorem tariffs fell by 4 percentage points and specific tariff rates fell by approximately 25% by the end of our sample.

Ultimately, our results suggest that there were two distinct liberalizations over the 1972-1988 period. Despite the absence of legislated tariff changes, the pre-Tokyo years saw substantial liberalization driven by price increases eroding the AVE of specific tariffs. This process continued throughout the 1980s, although lower inflation and the transition of some specific tariffs to ad valorem tariffs under the Tokyo Round mitigated the effect of this channel. Legislated tariffs, however, declined substantially throughout the 1980s, though this was masked at the aggregate level by changes in import composition.

Figure 5: Change in Legislated Tariff Rates 1972-1988



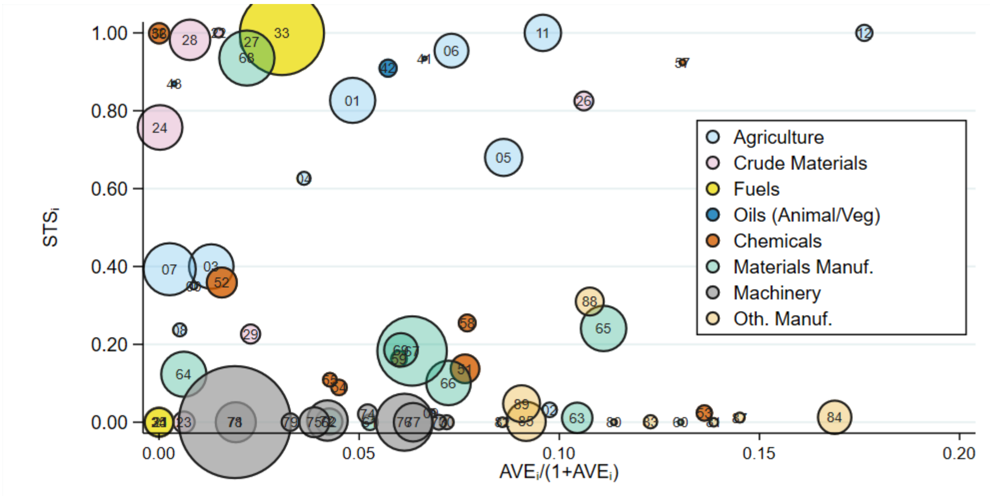
Notes: Figures plot point estimates and 95% confidence intervals from regressions of estimated column 1 legislated ad valorem tariff rates, $\ln(1+\hat{\tau}_{gt})$, in panel (a),$\ln \hat{f}_{gt}$ in panel (b) on time dummies and 5-digit TSUSA fixed effects. See main text for further details.

3.3 Industry Variation

Given the above results at the aggregate level, we briefly turn to cross-industry variation in tariff changes during our sample period. In particular we are interested in how industries differed in their exposure to the pre-1979 "accidental" liberalization and the post-Tokyo legislated liberalization.

We begin with the inflation-driven accidental liberalization. As discussed above, price increases lead to larger reductions in AVE tariff levels when specific tariffs account for a larger share of share of a variety's protection. One can show that this inflation exposure depends is the product of the specific tariff share (STS) of total protection and the proportion of tariffs in the tariff inclusive price of the good (AVE/(1+AVE)). The industries most exposed to the liberalization of the 1970s will thus be those that begin with higher AVE levels, specified primarily in per-unit terms. We identify such industries in Figure 6, which displays variation across 2-digit SITC sectors in AVE levels and in the share of tariff revenue coming from specific tariffs, both as of 1972. Sectors in the upper right of the figure – e.g., Tobacco (12), Plastics in Primary Form (57), Textile Fibers (26), Beverages (11), Sugars & Honey (6), Animal Fats & Oils (41) and Vegetable Fats & Oils (42)) – are those with both high AVE levels and substantial reliance on specific tariffs. It is these sectors that are exposed to liberalization in the presence of inflation, and thus drive the decline in aggregate AVE tariffs prior to 1980. Tariff liberalization in the 1970s, then, was primarily focused on agriculture and agriculture-adjacent industries.

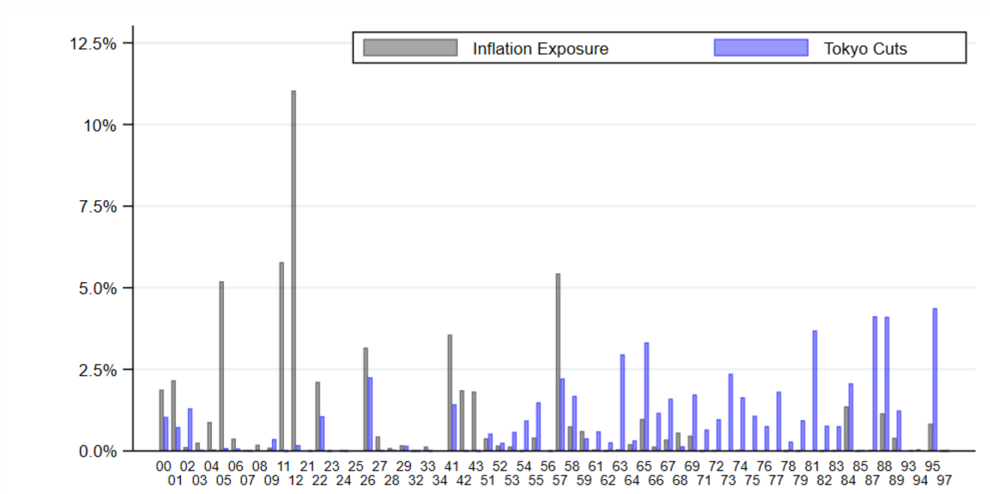
Figure 6: Sectoral Exposure in 1972 to Inflation Erosion of AVE (2-digit SITC)



Notes: Scatter plot of sectoral exposure to intra-policy tariff changes due to inflation/deflation as of 1972. Each bubble is a single 2-digit SITC revision 2 sector whose size reflects import share in 1972.

Conversely, in Figure 7 we illustrate exposure to the realized tariff reductions in the years before and after the agreement. The negotiated cuts tend to be more pervasive following the agreement, with the majority of industries experiencing tariff reductions.

Figure 7: Sectoral Exposure to Both Liberalizations (2-digit SITC)



4 Trade and 1980s Labor Market Outcomes

We have argued above that the effects of trade on the US economy prior to the 1990s remain understudied, and that the dataset we have constructed will prove valuable in addressing this shortcoming. To illustrate one application of these data, we revisit the literature analyzing the relationship between trade and inequality in the 1980s. This decade witnessed marked increases in inequality in terms of both wages and employment of skilled relative to non-skilled workers.[21] A sizable body of work has examined the role of trade in these changes, most often in the context of the Heckscher-Ohlin model (Slaughter, 2000). In this section, we combine our tariff estimates and a novel IV approach to estimate the effect of tariff liberalization on labor market outcomes between 1979 and 1988.

4.1 Identification

Any analysis of the effect of tariff liberalization on labor market outcomes must confront two fundamental identification challenges. First, the extent of tariff liberalization in different industries may have been chosen by policy makers with industry characteristics in mind. Second, the extent of tariff liberalization across industries may depend on pre-liberalization tariff levels that were themselves a function of industry characteristics. In each case, an endogeneity problem arises when the industry characteristics themselves directly affect labor market outcomes.

To deal with these challenges, we exploit the fact that Tokyo Round negotiations began from a starting point where product-level tariff cuts were given by the "Swiss Formula". The Swiss formula

$$AVE_{g,1987}^{Swiss}(Z, AVE_{gt_0}) = \frac{Z \times AVE_{gt_0}}{Z + AVE_{gt_0}} \tag{11}$$

determines the final post-Tokyo 1987 tariff AVE_g on good g solely as a function of two variables: (i) the initial product-level tariff AVE_{gt_0} on good g at time t_0 and (ii) a parameter Z that corresponds to the maximum final tariff rate across all products, which does not vary across goods. As such, for *given* pre-existing tariff levels AVE_{gt_0}, the single parameter Z determines the distribution of tariff liberalization across goods. Thus, by construction, the extent of liberalization is not correlated with industry characteristics conditional on pre-existing tariff levels. As a result, endogeneity concerns surrounding Tokyo Round tariff cuts center solely around the endogeneity of pre-existing product-level tariffs. To address this concern, we take advantage of the fact that "Column 2" tariffs are not re-negotiated in GATT rounds and were set by the Smoot-Hawley Tariff Act in 1930. While tariff levels in 1930 are unlikely to reflect the economic concerns of the 1980s, they are potentially

[21] For an excellent survey of this literature, see Helpman (2018).

correlated with NTR tariffs as of 1979, making them candidates for use in our instrument.

We combine the parameter Z with legislated Column 2 tariffs to construct an instrument for product-level tariffs AVE_{gt} as follows for the years 1979 to 1987.

$$AVE_{gt}^{IV} = AVE_{g,1978}^{Col2} + \frac{t-1979}{8}\left[AVE_{g,1987}^{Swiss}\left(Z, AVE_{g,1978}^{Col2}\right) - AVE_{g,1978}^{Col2}\right]$$

The instrumented value for AVE_{gt} in 1979, the year before the Tokyo tariff cuts begin, is the legislated Column 2 AVE tariff $AVE_{g,1979}^{Col2}$. The instrumented value in 1987 and also 1988, once the Tokyo tariff cuts have been fully implemented, is given by the value AVE_{g}^{Swiss} implied by Swiss formula in 11 using a value of $Z = 0.14$ (Deardorff and Stern, 1979). The instrumented values in the Tokyo phase out years are linearly interpolated between the pre-Tokyo and post-Tokyo instrumented values. Given the parameter Z does not vary across industries, the key exclusion restriction requires that omitted variables such as underlying industry characteristics do not jointly determine 1980s labor market outcomes and 1930 US column 2 tariffs.[22]

To construct the instrument, we digitize the US legislated column 2 tariffs from the fifth supplement to the 1978 TSUSA schedule. This is the final publication of the schedule prior to the Tokyo Round. We then apply these tariffs to all varieties in our data. This yields the observed tariff for country-varieties subject to column 2 rates. But, it reverts protection to the higher levels legislated in 1930 for countries receiving column I rates or duty free status. Using this process we calculate AVE tariffs as well as the Swiss formula instrument at the level of 4-digit SITC rev. 2 classification and then concord to the 4-digit SIC-level.

Two factors determine the instrument's strength. The first is the correlation between column 1 and column 2 tariffs. Similar to the PNTR literature, we find that this correlation is strong. In 1979, industries' observed AVE rates exhibit a 63% correlation with an analogously constructed rate under the column 2 legislated tariffs. The second factor governing the strength of our instrument is the extent to which the post-Tokyo tariff liberalization followed the Swiss formula in practice. To this end, Figure 8 plots the absolute value of Tokyo tariff cuts observed in the data against their Swiss-implied value. These measures are constructed at the tariff-line level for column 1 rates, holding prices fixed at their 1979 level. While there is deviation from the Swiss-implied liberalization, the overall strength of the relationship is strong: to a large extent, the US indeed liberalized as dictated by the Swiss formula. Specifically, the correlation between the Swiss-implied and observed tariff cuts in this figure is 0.73.

[22] This very closely resembles the exclusion restriction in the large literature exploring the causal effects of the US trade policy change to permanently normalize trade relations with China in the early 2000s. This change permanently imposed the US column 1 tariffs on China, eliminating uncertainty over whether Chinese exporters would face the US column 1 or column 2 tariffs.

Finally, a simple re-arrangement of the Swiss formula allows us to construct the coefficient Z implied by the observed changes in tariffs. To the extent that the formula was followed, we should observe a mass around a value of 0.14. We plot the distribution in Figure 9. While there is dispersion among goods, the median and modal coefficient was 0.14 as previously asserted. This again suggests a strong relationship between the liberalization implied by the Swiss formula and the one that took place.

Figure 8: Sectoral Exposure in 1979 to Tokyo Phaseouts

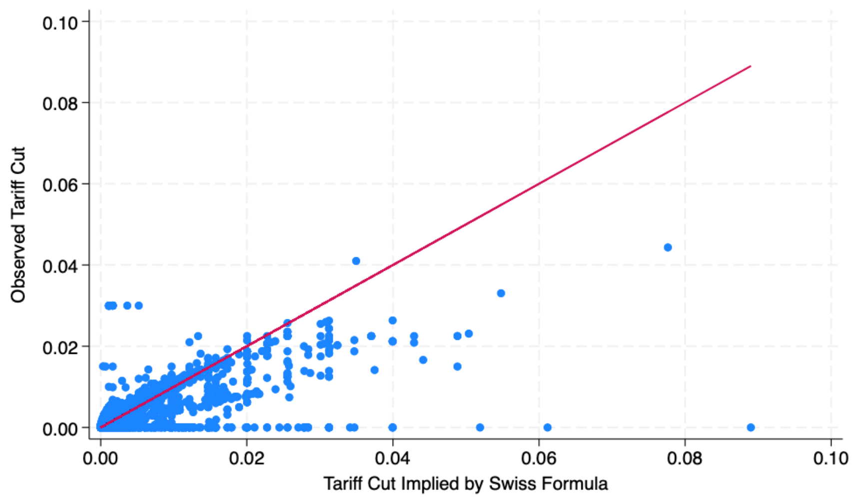
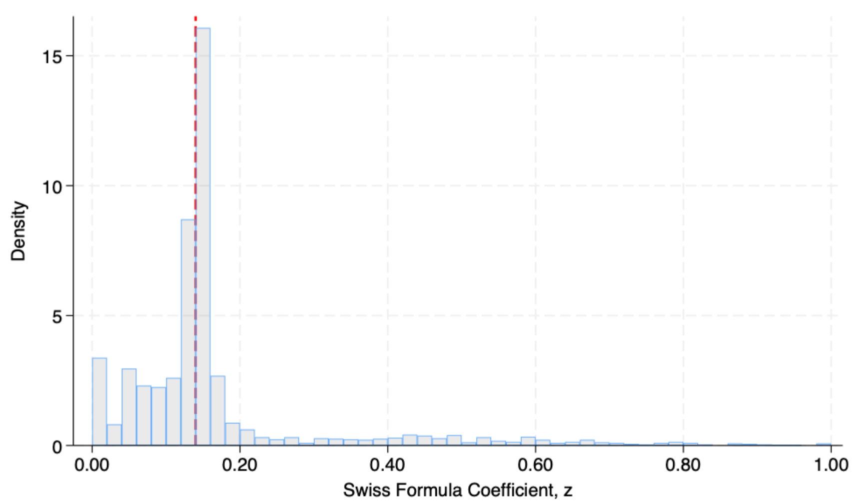


Figure 9: Distribution of Swiss Z Coefficient by TSUSA-5



4.2 Empirical Specification and Data

We now turn to the relationship between tariffs and industry-level outcomes in the US manufacturing sector between 1979 and 1988. At the four-digit SIC87 level, we consider the following specification:

$$y_{it} = \beta_0 + \beta_1 AVE_{it} + \gamma_i + \gamma_t + \epsilon_{it} \tag{12}$$

where y_{it} is an outcome of interest for industry i in year t and AVE_{it} is the four-digit SIC87 level AVE tariff in industry i and year t. The parameters γ_i and γ_t are industry and year fixed effects, respectively, while ϵ_{it} is the error term. We estimate (12) using both OLS and IV, employing our Swiss instrument from equation (4.1). We cluster the standard errors at the unit of tariff variation observation, which is industry-by-year.

As outcomes, we collect industry-level data from the NBER-CES Manufacturing Database (Becker et al., 2021). To explore how industries respond to the Tokyo tariff cuts at an aggregate level, we consider the real value of shipments as a proxy for output. To investigate distributional effects, we examine labor payroll as a share of the value of shipments and the ratio of labor payroll for non-production workers relative to production workers. The former measure captures any distributional effects between labor and other factors, while the latter will capture any distributional effects between skilled and unskilled workers as proxied, respectively, by non-production and production workers. Finally, to better understand how changes in labor payrolls combine effects on employment and wages, we also examine employment and wages for labor overall as well as for non-production workers relative to production workers.

4.3 Results

We begin by exploring the relationship between tariff liberalization and trade flows. In columns 1 and 2 of Table 2, we regress log industry imports and industry import penetration, respectively, against industry AVE using an OLS specification.[23] Column 1 shows no relationship between tariff levels and log imports, while column 2 suggests a small negative relationship between tariffs and import penetration. Specifically, column 2 implies that an interquartile increase in AVE tariffs correspond to a relative reduction in import penetration of 0.007, compared to a mean level of .07. Thus, an OLS specification suggests at most a relatively weak relationship between tariffs and imports following the Tokyo Round.

[23] Import penetration is defined as the ratio of imports to consumption, shipments and net imports where net imports is imports minus exports.

These two columns, however, ignore the concern that the distribution of industry-level tariff cuts may reflect factors that directly affect import volumes, such as demand or productivity shocks. To address this, columns 3 and 4 of the table replicate columns 1 and 2, instrumenting with the Swiss formula instrument defined in equation (4.1). As discussed above, the instrument is very strong with an F-statistic above 265. Moreover, accounting for potential endogeneity in tariff levels alters our results dramatically. Column 3 implies that a one percentage point reduction in AVE levels increases imports by .07 log points which is an order of magnitude larger than the OLS point estimate. Similarly, column 4 implies a response of import penetration to AVE levels more than five times as great as that implied by column 2. These columns suggest that appropriately accounting for potential endogeneity in tariff levels is of first-order importance: omitted variables correlated with both AVE tariffs and import volumes mask the relationship between tariff liberalization and import growth.

It is possible that domestic tariff reductions proxy for the effect of tariff liberalization more broadly. This would complicate interpretation of our estimates, as they capture some combination of import exposure and greater access to US export markets. To the extent that this is the case, we would expect AVE tariffs to be correlated with export flows as well as imports. Columns 5 and 6 of the table show that this is not the case. That is, reductions in instrumented AVE levels predict relative increases in industry imports, but not exports.

Table 2: Tariff Liberalization and Trade

	(1)	(2)	(3)	(4)	(5)	(6)
	$\ln(Imports)$	$ImportPenetration$	$\ln(Imports)$	$ImportPenetration$	$\ln(Exports)$	$\frac{Exports}{Shipments}$
AVE_{gt}	-0.692	-0.163***	-7.077***	-0.928***	-0.437	-0.0950
	(0.474)	(0.0463)	(0.865)	(0.0941)	(0.609)	(0.0807)
Esimation	OLS	OLS	IV	IV	IV	IV
1^{st} Stage Coeff.	-	-	.16	.16	.16	.16
1^{st} Stage F	-	-	272.16	270.65	267.57	270.65
Obs.	3693	3716	3693	3716	3689	3716

Notes: Estimating equation is equation (12) with dependent variable indicated in column header. Columns (1)-(3) estimated via OLS and columns (3)-(6) estimated using IV where AVE tariff instrumented using the Swiss IV defined in equation (4.1). All data measured or concorded to the 4-digit SIC87 classification. Standard errors clustered at the SIC87-year level. F-statistic is Kleibergen-Paap Rk F-statistic. *, **, and *** indicate p-values less than 10%, 5% and 1% respectively.

Given these results, we now turn to the effects of tariff liberalization on labor market outcomes. In columns 1 and 2 of Table 3, we examine the relationship between AVE levels and the log value of industry shipments and the share of that value accounted for by payments to labor, respectively. The columns suggest that reduced tariffs increased the value of shipments, potentially suggesting a role for increased offshoring (Feenstra and Hanson, 1999). Consistent with this, column 2 suggests that labor did not share evenly in the returns to increased output: an interquartile reduction in AVE levels corresponds to a reduction in labor share's of the value of shipments by 0.007, relative to a mean of 0.20. Thus, industries that received tariff cuts expanded, but the majority of the associated gains went to factors other than labor. In addition to the relative shift away from labor, column 3 suggests a further shift away from production workers in particular. Tariff reductions correspond to an increase non-production labor payroll relative to production labor payroll. These effects, too, are economically meaningful: a 4 percentage point reduction in AVE levels maps to a 0.035 log point increase in the relative payroll of non-production workers.

Table 3: Swiss IV and Labor Markets

	(1) $\ln(Shipments)$	(2) $\frac{LaborPay}{Shipments}$	(3) $\ln\left(\frac{Pay_{NP}}{Pay_P}\right)$	(4) $\ln\left(\frac{Emp_{NP}}{Emp_P}\right)$	(5) $\ln\left(\frac{Wage_{NP}}{Wage_P}\right)$
AVE_{gt}	-0.897**	0.163***	-0.846***	-0.941***	0.0955
	(0.380)	(0.0305)	(0.246)	(0.280)	(0.189)
1^{st} Stage Coeff.	.16	.16	.16	.16	.16
1^{st} Stage F	269.61	269.61	269.61	269.61	269.61
Obs.	3729	3729	3729	3729	3729

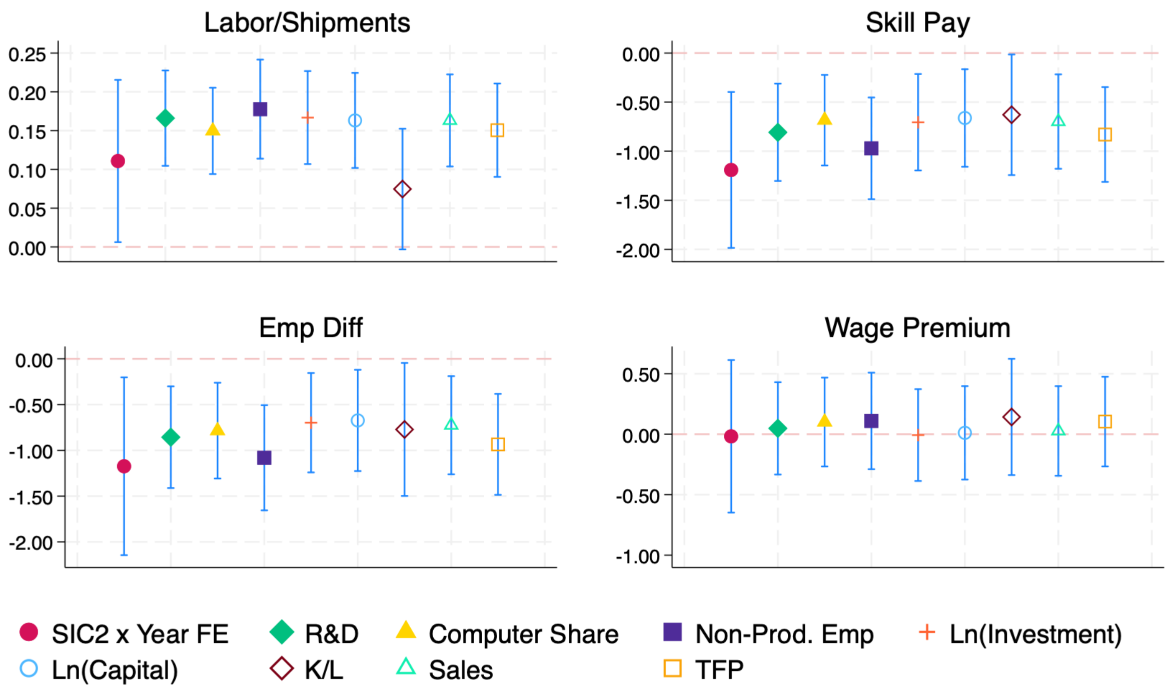
Notes: Estimating equation is equation (12) with dependent variable indicated in column header. Estimation via IV where AVE tariff instrumented using the Swiss IV defined in equation (4.1). All data measured or concorded to the 4-digit SIC87 classification. Standard errors clustered at the SIC87-year level. F-statistic is Kleibergen-Paap Rk F-statistic. *, **, and *** indicate p-values less than 10%, 5% and 1% respectively.

Columns 4 and 5 explores the distributional effects further by separating the effect on labor payments into its wage and employment components. The difference here is stark: the relative growth in payments to non-production workers is driven entirely by employment shifts, with no accompanying role for wages. Despite the difference in our approach, this is consistent with the limited role for trade in much of the empirical work of the 1990s, which focused largely on wage outcomes.

In Figure 10, we explore the robustness of these results to the inclusion of additional controls. Our primary concern is that column 2 tariffs, and thus our instrument, may be correlated with industry-level characteristics that simultaneously help explain changes in labor market outcomes in the post-Toyko years. Chief among these is industry susceptibility to skill-biased technological change. To address this, in Figure 10 reports results that repeat the specification from columns 2-5 of Table 3 but add each of the following industry-level variables interacted with a trend variable: R&D, the share of inputs sourced from the "computer" industry, non-production employment share, log investment, log capital stock, log capital-to-labor ratio, log sales, and TFP. Each of these variables is measured as of 1972, with the exception of R&D and computer input shares which are measured in 1975 and 1992, respectively. Finally, we also include two-digit SIC sector-by-year fixed effects. As is clear from the figures, our results are broadly robust to inclusion of these controls. That is, instrumented tariff levels are not simply capturing the effect of differential trends in pre-existing industry characteristics.

Our results suggest a shift away from both domestic labor and from production workers towards

Figure 10: Robustness



Notes: Each bar is the AVE tariff coefficient from a separate regression adding a single additional control to our baseline specification in equation (12). These controls are indicated in the legend. Estimation via IV where AVE tariff instrumented using the Swiss IV defined in equation (4.1). All data measured or concorded to the 4-digit SIC87 classification. Standard errors clustered at the SIC87-year level Bar whiskers indicate 95% confidence intervals.

non-production workers in response to tariff reductions following the Tokyo Round. This begs the question as to why. One possible explanation is that trade promoted innovation that was skill-biased in nature. To explore this possibility, we repeat the specification from Table 4, but take two measures of innovation as our outcome. First, we consider industry-level investment. Consistent with a story in which trade promotes innovation, we find that industries experiencing tariff reductions saw increased investment relative to less-exposed industries. Investment, of course, does not necessarily correspond to innovation. As an attempt to more directly capture innovation, we explore the relationship between industry level patent activity and AVE levels. Here, too, we find a positive response to reduced AVE levels. The point estimate in column 2 implies that an interquartile reduction in tariffs corresponds to a 0.05 log point increase in industry patenting.

Table 4: Swiss IV and Innovation

	(1) ln ($Invest$)	(2) ln ($Patent$)
AVE_{gt}	-1.985***	-1.243***
	(0.698)	(0.403)
1^{st} Stage Coeff.	.16	.15
1^{st} Stage F	269.61	247.32
Obs.	3729	3517

Notes: Dependent variable indicated in column header. All data are measured or concorded to the 4-digit SIC 1987 classification. Standard errors are clustered at the sic by year level. Independent variable is AVE tariff level which is instrumented based on the column II tariffs and the swiss formula. All specifications include time and year fixed effects. *, **, and *** indicate p-values less than 10%, 5% and 1% respectively.

5 Conclusion

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

With this new tariff-line level data, we explore the evolution of US tariffs in this era. Importantly, specific tariffs account for between 25%-35% of US tariff lines during this era. While typically overlooked, we show that specific tariffs play a critical role in shaping the evolution of aggregate ad valorem tariff rates in this era. Rampant inflation eroded the protective capacity of specific tariffs and played as large a role in reducing tariff protection as the negotiated tariff reductions themselves.

The formulaic nature of negotiations over tariff cuts under the Tokyo round of the GATT allows construction of a simple IV that we leverage to circumvent concerns about endogenous changes in tariff protection. We use this instrument to quantify the effect of tariff changes on industry outcomes from 1979-1988. We show that tariff cuts lead to rising import penetration and were associated with a decline in the labor share. Moreover, these tariff cuts lead to rising income inequality rises between skilled and unskilled workers that is driven by rising skill-intensity and not rising wage inequality. Combined with the additional results that tariff cuts stimulated investment and patenting, our analysis points toward a prominent role in this era for trade-induced skill-biased technological change.

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

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