Trade and US Inequality in the Tokyo Round*

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Abstract

Against a backdrop of sharply rising inequality, the Tokyo Round of the GATT resulted in a 1.6 percentage point reduction in average US tariffs – larger than CUSFTA, NAFTA, and the liberalization accompanying the granting of PNTR to China. We construct a novel IV based on the so-called "Swiss formula" that governed the Tokyo Round tariff liberalization to provide evidence of its effects on imports and inequality. Instrumented tariff reductions explain approximately 20% of the rise in income inequality between non-production and production workers between 1979 and 1988. This effect is largest among women, workers in routine occupations, and workers in more technology-intensive industries, suggesting a complementarity between trade liberalization and skill-biased technological change.

JEL codes: F13, F14, F66

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1 Introduction

Over the course of more than a decade, an enormous body of work has documented large and persistent consequences of trade for US firms and workers, exploiting differential exposure across geographic locations (Autor et al., 2013), industries (Pierce and Schott, 2016), or some combination of the two (Hakobyan and McLaren, 2016). In doing so, this literature has overwhelmingly focused on the post-1990 era of "hyper-globalization." This focus, however, comes with an important but often overlooked caveat: by the 1990s, much of the post-war US trade liberalization process was already complete. The US ad valorem equivalent (AVE) tariff, defined as the ratio of duties collected to total imports, stood at approximately 20% following the Tariff Act of 1930 – the so-called "Smoot-Hawley" tariff. By 1990, this had fallen to 3.3%.[1] This suggests that much of the economic impact of trade on the US economy was likely felt in the decades prior to those emphasized by scholars.

In particular, the distributional effects of the Toyko Round of the General Agreement on Tariffs and Trade (GATT) – a multilateral liberalization that governed US tariff reductions between 1980 and 1987 – have received almost no attention. This is despite the fact that, in terms of the resulting aggregate tariff changes, the Tokyo Round represents a larger liberalization than the Canada-US Free Trade Agreement (CUSFTA) in 1989, the North American Free Trade Agreement (NAFTA) in 1994, or the granting of Permanent Normal Trade Relations (PNTR) to China in 2001. Between 1979 and 1987, the average tariff across US manufacturing imports fell by 2.2 percentage points, resulting in a 1.4 percentage point decline in the US average tariff. The corresponding reductions in the US average tariff are 1.07 percentage points under PNTR, 0.75 percentage points under CUSFTA, and 0.14 percentage points under NAFTA.[2] Indeed, from this perspective, the Tokyo Round tariff cuts were nearly three-quarters as large as these three major shocks combined.[3]

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

[2] The average tariff decline across manufacturing and non-manufacturing imports was 4 percentage points for Canada under CUSFTA (Trefler, 2004), which at its onset in 1989 applied to 18.8% of US imports ($-4 \times 0.188 = -0.75$ percentage points.) Similarly, a 2.1 percentage point cut for Mexico under NAFTA (Hakobyan and McLaren, 2016) impacted only 6.7% of imports in 1993. And while tariffs on Chinese imports did not change following PNTR, Handley and Limão (2017) estimate a reduction in uncertainty equivalent to a 13 percentage point reduction in tariffs, covering 8.2% of US imports in 2000 (USITC Dataweb). The manufacturing share of 1978 US imports was 63.5%, with the Tokyo Round liberalization covering more than 95% of all 1978 US imports and approximately 75% of these imports facing strictly positive tariffs.

[3] One potential reason that prior work has overlooked the Tokyo Round is that, as we detail in Greenland et al. (2025), the substantial tariff cuts in this period were accompanied by a compositional shift in US imports towards high tariff goods that left the import-weighted average tariff relatively constant.

In this paper, we explore the effects of the Tokyo Round liberalization on inequality between production and non-production workers in the US manufacturing sector. Our approach exploits newly available data on US tariffs in the era (Greenland et al., 2025) and a unique feature of the Tokyo Round negotiations that allows us to circumvent standard concerns regarding endogenous tariff policy. We instrument for US tariff changes between 1979 and 1988, exploiting the fact that the so-called "Swiss formula" served as the starting point for Tokyo Round negotiations. The Swiss formula compressed the tariff distribution by assigning larger tariff reductions to goods with higher pre-Tokyo tariff levels. Crucially, the only reason for differential liberalization across goods under the Swiss formula was differential pre-Tokyo tariffs.[4] Instrumenting for observed tariff changes with those dictated by the Swiss formula thus allows us to avoid the concern that US policymakers may have protected certain industries in this period due to, e.g., expected import growth.

To address the concern that pre-Tokyo tariffs reflect historical protection choices of policymakers, our instrument replaces pre-Tokyo column 1 (MFN) tariffs in the Swiss formula with their column 2 counterpart. Column 2 tariffs were primarily determined by the 1930 Smoot-Hawley tariffs (Pierce and Schott, 2016; Handley and Limão, 2017) and were reserved for a small subset of communist countries at the onset of the Tokyo Round. The exclusion restriction for our "Swiss IV" thus rests on the 1930 tariffs being uncorrelated with unobserved factors that drive changes in 1980s labor market outcomes. While the instrument is correlated with a host of 1978 industry characteristics, we show this correlation is driven by persistence in industry characteristics from the Smoot-Hawley era. As policymakers in 1930 were unlikely to set tariffs based on anticipation of 1980s economic shocks such as skill-biased technological change, this mitigates the concern that our instrument proxies for important unobservables. For example, conditional on 1929 industry characteristics, we show that our instrument is systematically unrelated to proxies for 1980s skill-biased technological change. Ultimately, we argue that our Swiss IV has strong claims to reflecting exogenous tariff cuts.

We first use this instrument to estimate the impact of Tokyo Round tariff cuts on US industry-level trade flows between 1979 and 1988. We find a substantial impact on imports and our estimated trade elasticity of 9 falls within the typical range for similar empirical

[4] Our approach is part of a large body of work in which the extent of liberalization depends on pre-liberalization tariff levels. See work on, e.g., Chile (Pavcnik, 2002), Colombia (Goldberg and Pavcnik, 2005), India (Topalova, 2010; Topalova and Khandelwal, 2011), CUSFTA (Romalis, 2007; Kovak and Morrow, 2022), NAFTA (Caliendo and Parro, 2015; Hakobyan and McLaren, 2016; Besedes et al., 2020; Benguria, 2023), and the US granting normal trade relations (NTR) status to Vietnam (McCaig and Pavcnik, 2018) and PNTR to China (Pierce and Schott, 2016; Handley and Limão, 2017; Greenland et al., 2019).

settings (Ruhl, 2008; Head and Mayer, 2014). Exports, on the other hand, are not responsive to changes in US tariffs, which suggests that our results are not primarily driven by simultaneous Tokyo Round tariff cuts by US trade partners. These results are robust to a wide range of controls for alternative trade policy shocks in the era.

The instrument also reveals an important role for tariff changes in driving within-industry pay inequality between non-production and production workers.[5] As in the earlier trade and technology literature (Lawrence and Slaughter, 1993; Berman et al., 1994; Sachs et al., 1994; Berman et al., 1998), we define industry-level pay inequality as non-production worker pay relative to production worker pay, where industry-level pay for each group is the product of employment and wages. We find that an interquartile increase in the magnitude of the industry-level tariff liberalization throughout our sample period increases log pay inequality by approximately 0.08. In terms of economic significance, we show that import tariff liberalization can account for roughly 20% of the observed increase in log pay inequality in our sample period, with all of this occurring within as opposed to between industries. Decomposing this effect into changes in the relative employment of non-production to production workers and changes in relative wages between worker types, we show that the increase in pay inequality is primarily driven by the employment margin.

Exploring heterogeneity across workers, occupations, and industries, we document a striking pattern: the increase in pay inequality in response to the Tokyo Round is driven overwhelmingly by women. Indeed, the magnitude of the effect of falling tariffs on rising inequality among women is approximately double that of the manufacturing sector as a whole. We show that this is due in part to the fact that rising inequality within manufacturing is most pronounced in routine occupations, which account for a disproportionate share of women's employment in the era. Turning to the role of technological change, we find evidence that trade and technology interact to shape inequality in this period. In particular, we show that technology-intensive industries, as captured by higher pre-existing levels of automation, drive the increasing income inequality that accompanies the Tokyo Round liberalization.

While we see the paper as an important case study of the Tokyo Round, it also complements the large literature on 1980s trade and inequality more generally. Early work in this literature struggled to identify trade-driven changes in goods prices.[6] Furthermore, much

[5]Throughout the paper, we refer to income and pay interchangeably.

[6]See Slaughter (2000) and Feenstra and Hanson (2003) for excellent surveys. Slaughter (2000) in particular notes that this literature "made substantial progress understanding how to relate a given change in relative product prices to changes in relative factor prices ... [but] made less progress understanding whether these product-price changes have any thing to do with international trade."

of this literature focused on inequality between industries, often through the lens of factor endowment theory via the Heckscher-Ohlin model. In contrast, the broader labor literature documented that most changes in skill intensity happened within industries and especially in occupations and industries with more scope for adopting new skill-biased technologies (Berman et al., 1994, 1998; Acemoglu and Autor, 2011). Thus, the early literature concluded that trade had little effect on inequality (Berman et al., 1994, 1998). Conversely, our findings reveal a crucial role for trade liberalization in driving increased inequality within industries, especially in occupations and industries with more scope for adopting new skill-biased technologies.

More recently, Batistich and Bond (2023) identify notable effects on racial inequality in US labor markets due to rising imports from Japan in the 1970s and 1980s. In particular, they find relative reductions in manufacturing employment and labor force participation and declining wages for black males compared to white males. We differ from these prior studies by focusing on a broad-based, multilateral policy shock with an approach that does not rely on assumptions about the exogeneity of export growth in foreign markets, but rather on the institutional details of historical US tariff policy.[7] While complementary to existing work, our results emphasize an important role for trade policy in explaining inequality. The paper also contributes to the numerous studies exploring the relationship between trade and technology. Consistent with the disparate theoretical predictions on the link between import competition and innovation (Lawrence, 2000; Acemoglu, 2003; Aghion et al., 2005), empirical evidence on this relationship is mixed.[8] We argue that trade and technology jointly, rather than separately, drive inequality.

A small number of papers have explored the effects of trade policy in this era. In Greenland et al. (2025), we use annual US legislated tariff data between 1972 and 1988 to show that the protection provided by specific tariffs was eroded by 1970s inflation despite the absence of tariff policy changes, and that this "accidental liberalization" was larger in magnitude than the Tokyo Round legislated tariff cuts. Most other work exploring the Tokyo Round took place during the early years of its phase-in and focused on predicted effects. Several papers emphasize consequences for the structure of production across member countries using

[7]In separate work, Lu and Ng (2013) instrument for US import penetration with UK import penetration and find that US manufacturing industries between 1971 and 2001 respond to import competition by substituting away from routine occupations towards non-routine occupations. This requires the assumption that UK import penetration is driven by supply-side forces in the rest of the world and not, for instance, UK demand shocks that are correlated with those in the US. Our approach requires no such assumption.

[8]For instance, Bustos (2011), Mion and Zhu (2013) and Bloom et al. (2016) find positive effects of trade on innovation, while Autor et al. (2020) and Aghion et al. (2024) find negative effects.

general equilibrium models (Brown and Whalley, 1980; Deardorff and Stern, 1981, 1983), while others focus on the nature of the negotiations themselves (Ahmad, 1978; Chan, 1985). Closer to us, Gaston and Trefler (1994) explore the labor market consequences of tariffs in the period. Using cross-industry variation in 1983 US tariffs, they find that more protected industries paid higher wages but find no differential impact on imports. Moreover, they argue that policymakers did not manipulate the Tokyo Round tariffs in ways that affected labor market outcomes or imports. In contrast, our Swiss IV suggests a substantial role for endogenous policy setting.

The paper proceeds as follows. Section 2 describes the data. Section 3 describes our estimation strategy and identification challenges. Section 4 describes baseline results and robustness, while Section 5 explores heterogeneity in these responses. Section 6 concludes.

2 Data

To evaluate the effects of the Tokyo Round on inequality, we require information on industry outcomes, tariffs, and trade flows. Our final dataset draws from a number of main sources. Primary industry outcomes come from the NBER-CES Manufacturing Industry Database (Becker et al., 2021) and from IPUMS (Ruggles et al., 2024). From these datasets we collect information on wages and employment at the industry level separately for production and non-production workers.[9] IPUMS data also allow us to separate workers by sex and occupation. The NBER data are available annually, while IPUMS data are reported for each census year. Annual industry imports and exports come from Schott (2008). The NBER and trade flow data measure industry outcomes at the 4-digit level of the 1987 Standard Industrial Classification (SIC) system, while IPUMS data are available at the Census industry level. Finally, we take tariff line-level legislated US tariffs from Greenland et al. (2025) and, as Appendix B.1 describes in detail, concord them to calculate AVE tariffs at the industry level.[10] Appendix Table A.1 contains summary statistics.

Figure 1 about here.

Before moving to our empirical methodology in Section 3, we briefly describe the trade

[9] We use "non-production" and "production" to be consistent with the definitions in the NBER-CES data. Scholars from the era alternatively referred to these groups as "skilled" and "unskilled", respectively.

[10] Because the US uses both ad valorem and specific tariffs in this era, the dataset contains legislated ad valorem tariffs and the AVE of legislated specific tariffs that together sum to the AVE legislated tariff for the good. The AVE of a legislated specific tariff is the legislated specific tariff divided by its unit value. The dataset covers over 5000 5-digit TSUSA goods and over 98% of imports during our sample.

environment as well as the evolution of inequality during our sample. To this end, Figure 1 presents a heat map describing 1980s US import growth across sectors and partner countries. Descending along the vertical axis are the largest exporters to the US in 1978, with Canada and Japan comprising more than 40% of US imports immediately preceding the Tokyo Round. The horizontal axis orders 2-digit SIC sectors from largest to smallest in terms of the sector's 1978 share of US imports. Cooler colored squares in the figure represent smaller log import growth between 1979 and 1988, while warmer colors represent larger growth.

The key takeaway from Figure 1 is that, unlike many other episodes studied in the literature on trade and labor in the US, import growth following the Tokyo Round is not limited to a few countries nor to a few sectors. Illustrating the breadth of import growth across countries, log imports from every country except Germany grow by at least 1.25 (250%) in at least two sectors. Similarly, log imports in every sector except agriculture see growth of at least 1.25 from at least two countries. Moreover, import growth is negative only for Japanese primary metals manufacturing and Hong Kong petroleum. That is, while import growth is most pronounced in high-tech sectors such as electronics equipment, computer equipment, and measuring instruments and from countries that were not the top few US trade partners – i.e., countries other than Canada, Japan, and Germany – import growth is not limited to particular trade partners or industries.

Figure 2 about here.

This broad-based growth in imports mirrors the nature of the tariff liberalization following the Tokyo Round. Figure 2 displays AVE tariffs in 1978, with each dot representing a 4-digit SIC industry, and with industries grouped according to their 2-digit SIC sector. On average, 4-digit industries have a pre-Tokyo AVE tariff of 6.7%. However, the variation is considerable, both within and across sectors, with a standard deviation of 5.4%. The variation in pre-Tokyo AVE tariffs corresponds directly to variation in Tokyo Round tariff cuts since, as we discuss in detail below, the Swiss formula reduces tariffs by more for industries with higher pre-Tokyo AVE tariffs. In sum, the Tokyo Round led to large and varying exposure to tariff liberalization both within and across sectors.

Figure 3 about here.

Finally, we display the industry-level variation in changes in pay inequality between non-production and production workers throughout our sample. Because industry-level pay for each group of workers is the product of employment and wages, the change in log pay inequality is the sum of the changes in log employment and log wages of non-production workers relative to production workers. Figure 3 displays kernel densities of these log changes between 1979 and 1988. The gray distribution outlined in black depicts industry-level changes in pay inequality. The blue and red distributions represent the two components of pay inequality: employment of non-production relative to production workers and the wage of non-production relative to production workers, respectively. Two points stand out in the figure. First, more than 80% of industries see rising income inequality, producing an average increase of approximately 0.17 log points. Second, this is accompanied by rising non-production employment intensity which grows, on average, by 0.16 log points. In contrast, on average, wage inequality increases by a modest 0.01 log points, with a distribution that is nearly symmetric around 0. In formally exploring the effects of the Tokyo Round liberalization along each of these margins, we find a similarly dominant role for the employment channel in driving income inequality. We return to this below.

3 Empirical Methodology

We explore the effects of tariff liberalization under the Tokyo Round on US trade flows and industry outcomes. The sample period for our outcome variables in the trade and NBER-CES manufacturing data begins in 1979 – the year before the US began implementing Tokyo Round tariff cuts – and ends in 1988 – the year after the US completed implementing the liberalization. In our IPUMS specifications, we examine changes between 1980 and 1990. In both settings, we estimate industry-level first-difference regressions of the form

$$\Delta y_i = \beta_0 \Delta \text{ln}(1 + AVE_i) + \mathbf{X}_i \boldsymbol{\beta} + \varepsilon_i \tag{1}$$

where, for industry i, Δy_i is the change in an outcome of interest between the first and last year in the sample and $\Delta \ln(1 + AVE_i)$ is the log change in one plus the industry's AVE tariff. $\mathbf{X}_i$ is a vector of industry-specific controls and ε_i is the error term.

Equation (1) faces the standard identification concern that tariffs are set non-randomly. The specific concern in our context is that industry-level tariff reductions may be correlated with unobserved industry-level trends in import growth or labor market outcomes. For example, policymakers may choose to protect industries facing high expected import growth by assigning them smaller tariff reductions. Such endogenous policy setting would bias a negative coefficient of tariff changes on imports upwards towards zero, as larger tariff cuts would be associated with lower average import growth. Alternatively, against a backdrop of rising inequality and surging import competition, policymakers may choose to protect industries in which they expect rising inequality to be most pronounced. This would again bias any negative estimate of the relationship between tariff changes and changing pay inequality upwards towards zero, as larger tariff cuts would be associated with lower growth in inequality.

To address these concerns, we rely on a unique institutional detail of the Tokyo Round negotiations. In part due to fears that strategic protectionism would hamper the liberalization process, parties to the Tokyo Round sought a formula-based approach similar to the one taken under the Kennedy Round, which had aimed to reduce tariffs by 50%. While the US proposed a 60% reduction on all industrial tariffs under the Tokyo Round, the European Economic Community (EEC) preferred "harmonization", in which initially higher tariffs were reduced by more (Neu, 1979). Consistent with this aim, the Swiss delegation to the negotiations proposed a simple formula for post-Tokyo tariffs that depended only on a good's initial tariff. This so-called "Swiss formula" compressed the distribution of final AVE tariffs in year T by imposing a maximum AVE tariff of 0.14 and assigning larger tariff cuts to products with larger tariffs in an initial year t_0 (Swiss Delegation, 1976).[11] Specifically, the Swiss formula is

$$AVE_{iT}^{Swiss} = \frac{0.14 \times AVE_{it_0}}{0.14 + AVE_{it_0}}. \tag{2}$$

The extent of an industry's liberalization under the formula is thus completely determined by initial tariffs, AVE_{it_0}, which are the only potential source of endogeneity in the tariff liberalization implied by the formula. Importantly for our identification, Neu (1979, p.11) notes that "there was no historical or intellectual reason for the choice of this particular formula, it was chosen principally because it was simple and implied an acceptable average depth of cut." This suggests that the formula itself was not the product of political economy considerations related to, e.g., US imports and labor market outcomes.

That said, the extent to which countries followed the Swiss formula varied both across countries and across sectors within countries. While the US followed the Swiss formula much more closely than other countries, they did not strictly adhere to the formula (see, e.g. Table 14 of Deardorff and Stern (1979) or Table 6 of Neu (1979)). The US identified sectors,

[11] The tariff cut implied by the Swiss formula tariff is $AVE_{iT}^{Swiss} - AVE_{it_0} = \frac{-AVE_{it_0}^2}{0.14+AVE_{it_0}}$ and increasing in AVE_{it_0}. The maximum final tariff is $\lim_{AVE_{it_0}\to\infty} \frac{0.14\times AVE_{it_0}}{0.14+AVE_{it_0}} = \lim_{AVE_{it_0}\to\infty} \frac{0.14}{\frac{0.14}{AVE_{it_0}}+1} = 0.14$. The Swiss delegation proposed the value 0.14 and the US used this value (Swiss Delegation, 1976; Deardorff and Stern, 1979).

including textiles, apparel, leather, and footwear, as needing protection and these sectors indeed faced less than Swiss-implied tariff cuts. Other US sectors such as crops, tobacco, and chemicals faced less than Swiss-implied tariff cuts, likely due to other countries imposing less than Swiss-implied tariff cuts (Neu, 1979).

Figure 4 about here.

We explore the ultimate results of this process in Figure 4, which plots the 4-digit SIC AVE tariff change between 1979 and 1988 implied by the Swiss formula on the horizontal axis against the observed AVE tariff change over this period on the vertical axis.[12] The gray 45 degree line represents exact adherence to the Swiss formula, such that industries above the line are "under-liberalized" and receive less-than-prescribed tariff cuts while industries below the line are "over-liberalized" and receive more-than-prescribed cuts. Three points stand out in the figure. First, as discussed above, Tokyo Round tariff changes constitute a substantial liberalization: the mean observed tariff reduction is 2.2 percentage points, with a standard deviation of 2.3 percentage points. Second, the strong positive correlation of 0.57 between observed and prescribed tariff cuts implies that the Swiss formula does indeed serve as a benchmark for tariff liberalization under the Tokyo Round. Finally, the dispersion across industries in observed tariff cuts for a *given* Swiss-prescribed tariff cut – that is, the extent of the deviation from the 45 degree line – is an empirical representation of the concerns regarding endogenous tariff setting that motivate our instrument. In Appendix C, we explore the extent of under- or over-liberalization in more detail and show that industries with greater production shares of labor tended to be under-liberalized.

While tariff cuts under the Swiss formula are determined entirely by initial tariffs and a single known parameter, these initial tariffs are themselves non-random. In particular, the so-called column 1 (i.e. MFN) tariffs immediately preceding the Tokyo Round reflect the six GATT rounds between 1930 and 1979 and, in turn, the decisions of policymakers to protect certain industries during these rounds. For example, as we show in Appendix C, policymakers during the Tokyo Round tended to protect industries that used production labor more intensively from tariff cuts prescribed by the Swiss formula. As such, using the tariff variation implied by the Swiss formula may still generate biased estimates through the endogenous relationship between outcomes and initial AVE tariffs.

To address this concern, we rely on a second institutional feature of the US tariff schedule: while the US imposed column 1 tariffs on imports from most countries, it specified "column 2"

[12]See Appendix Figure A.1 for an analogous version of Figure 4 at the 5-digit TSUSA tariff line level.

tariffs that applied to a small subset of non-GATT countries.[13] These tariffs were themselves originally the MFN tariffs from the Smoot-Hawley Tariff Act of 1930. Indeed, Appendix Figure A.2 shows the strong correlation between 1930 Smoot-Hawley AVE tariffs and 1978 column 2 AVE tariffs.[14] As column 2 tariffs are based on 1930 MFN tariffs, there is a strong positive correlation between the 1978 column 2 tariffs and their column 1 counterparts. But as they are largely determined by tariffs nearly 50 years prior to our period of study, the Tokyo-era column 2 tariffs are less likely to have been set with an eye towards 1980s outcomes of interest than the 1978 column 1 tariffs.[15]

The construction of our instrument rests on each of the two institutional features of US tariff policy described above: that Tokyo Round negotiations used the Swiss formula as a benchmark and that column 2 tariffs were determined by the 1930 Smoot-Hawley tariffs but were still strongly correlated with their column 1 counterparts on the eve of the Tokyo Round. Specifically, we rely on the Swiss formula to dictate the magnitude of tariff cuts across goods g but substitute column 2 AVE tariffs into the formula in lieu of the observed column 1 AVE tariffs. At the 5-digit TSUSA level, we construct time-varying values of the Swiss measure. The value in 1988, after the phase-out has finished, is given by

$$AVE^{Swiss}_{g,1988} = \frac{0.14 \times AVE^{Col2}_{g,1978}}{0.14 + AVE^{Col2}_{g,1978}}. \tag{3}$$

The value in 1979, prior to the liberalization, is the 1978 column 2 AVE tariff, $AVE^{Col2}_{g,1978}$.[16] We aggregate $AVE^{Swiss}_{g,1988}$ to 4-digit SIC industries i, as detailed in Appendix B.1, and define

[13] As is well known, GATT Article I requires that GATT members apply MFN tariffs on imports from all member countries. This, however, leaves discretion for the tariffs imposed on non-members. The extent of discretion is substantial at the onset of the Tokyo Round, when fewer than 90 countries belonged to the GATT. The US imposed column 1 tariffs on the vast majority of non-GATT members during our sample period. But, it imposed column 2 tariffs on a small subset of communist non-GATT members, and even violated GATT Article I by imposing column 2 tariffs on some communist GATT members.

[14] The raw correlation is 0.71 and rises to 0.78 when weighted by 1979 imports. Products have, of course, been added to the tariff schedule since 1930. New products are typically classified by the Committee for Statistical Annotation of Tariff Schedules, more commonly known as the 484(f) Committee, by assigning them into existing tariff categories, including general "other" classifications. (https://www.usitc.gov/harmonized_tariff_information/modifications_to_hts).

[15] Appendix Figure A.3 illustrates the strong positive correlation between the column 1 and column 2 tariffs in 1978 at the 4-digit SIC industry level (panel (a)) and tariff-line TSUSA level (panel (b)). It also shows substantial dispersion across column 1 tariffs for a given column 2 tariff, implying that policymakers did indeed protect some industries more than others during the six preceding GATT rounds.

[16] We construct the instrument using the AVE in 1978 instead of 1979 to avoid concerns regarding endogenous determination of contemporaneous unit values. Our results are unchanged by using 1979 AVE tariffs.

our instrument as the change in the log of one plus the constructed Swiss measure:

$$\Delta \ln(1 + AVE_i^{IV}) \equiv \ln(1 + AVE_{i,1988}^{Swiss}) - \ln(1 + AVE_{i,1978}^{Col2}) \leq 0. \tag{4}$$

Figure 5 about here.

Figure 5 illustrates the two-stage least squares intuition behind our Swiss IV at the 4-digit SIC level. Panel (a) plots the first stage relationship: the instrumented change in AVE tariffs between 1979 and 1988 against the observed change in tariffs over this period.[17] Our instrumented change in AVE tariffs strongly predicts observed tariff changes. Industries with higher 1978 column 2 AVE tariffs, and thus greater predicted cuts by the Swiss formula, see substantially larger AVE tariff cuts following the Tokyo Round. Panel (b) displays the second stage relationship: predicted AVE tariff changes from panel (a) against observed import growth. The figure suggests that the Tokyo Round liberalization generated substantial import growth, as industries with larger predicted tariff reductions exhibit stronger import growth between 1979 and 1988. This relationship also suggests a role for Tokyo Round tariff cuts in explaining other industry outcomes in this era.

The exclusion restriction for our instrument is that unobserved factors that directly affect 1980s US trade or labor market outcomes are uncorrelated with our instrumented measure of Swiss-implied tariff liberalization. As the cross-industry variation in the instrument stems entirely from variation in 1978 column 2 tariffs, this amounts to a requirement that unobserved correlates of these tariffs do not determine 1980s changes in trade or inequality. To explore this, we conduct a number of balance tests. In particular, we regress $\Delta \ln(1+AVE_i^{IV})$ against three categories of industry-level characteristics: other trade policy related variables, measures of industry factor intensity, and measures of exposure to skill-biased technological change.

We include four trade-policy-related variables. To account for changes in export opportunities as a result of the Tokyo Round liberalization, we include an export-weighted average of changes in tariffs between 1979 and 1987 in Canada, Japan, and the EEC.[18] We also

[17] A small number of industries see AVE tariff increases. This is possible even in the absence of legislated tariff increases due to falling unit values in the presence of specific tariffs. Approximately one-third of US tariffs lines before the Tokyo Round faced specific tariffs (Greenland et al., 2025).

[18] For Japan and the EEC, we digitize tariff schedules from the WTO website: `https://www.wto.org/english/docs_e/gattbilaterals_e/indexbyround_e.htm`. We manually concord tariff lines to the Customs Co-operation Council Nomenclature (CCCN) and map these to 1992 HS codes using a concordance provided by WITS: `https://wits.worldbank.org/product_concordance.html`. Finally, we concord these to the SIC using the concordance created by Pierce and Schott (2012). We take Canadian import tariffs

include log changes in industry unit values between 1972 and 1979 and the share of the tariff-inclusive price in 1979 accounted for by specific tariffs.[19] These jointly capture exposure to inflation-driven changes in industry-level protection in the presence of specific tariffs (Greenland et al., 2025). Finally, we include a dummy variable equal to 1 if the industry was covered by the Multifiber Arrangement, which imposed import quotas on textiles and apparel (Harrigan and Barrows, 2009; Khandelwal et al., 2013).

We use various proxies for industry factor intensity and skill-biased technological change. Specifically, we include four proxies for industry factor intensity: the log capital-to-labor ratio, the share of employment accounted for by non-production workers, and the ratio of materials to shipments, each as of 1978, as well as the share of employment accounted for by women as of 1980. For skill-biased technological change, we include four proxies: log 1978 investment, the change in log investment between 1970 and 1978, the routine task-intensity index of Autor and Dorn (2013), and changes in automation between 1947 and 1978, following Acemoglu and Restrepo (2019, 2020).[20]

Table 1 about here.

In Table 1, we regress our instrument against the trade-policy-related variables in column 1, factor intensity measures in column 2, technological change measures in column 3, and the three sets simultaneously in column 4. The table clearly shows that our Swiss IV, and hence 1978 column 2 tariffs, is indeed correlated with a host of industry-level trade policy variables and industry characteristics.[21] Columns 1 and 4 emphasize the correlation between

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from the Canadian SIC system and concord these to the US SIC system using data and a concordance from Trefler (2004). We chose these countries due to their importance as US export destinations as displayed in Appendix Figure B.1. Appendix Figure B.2 scatters industry-level tariff changes by the US and foreign countries.

[19]The share of the tariff inclusive price accounted for by specific tariffs can be written as $STS_i \frac{AVE_i}{1+AVE_i}$, where STS_i is the share of tariff revenue in industry i generated by specific tariffs.

[20]We use data from the NBER-CES Manufacturing Industry Database (Becker et al., 2021) for capital-to-labor ratios, non-production employment shares, materials to shipments, and investment. Data on women's employment shares are calculated based on data from IPUMS (Ruggles et al., 2024), which we concord from the Census industry to the SIC system. The routineness measure of Autor and Dorn (2013) is calculated based on data from the Dictionary of Occupational Titles as of 1977 and defined at the occupation level - specifically, at the *occ1990dd* classification – as the log of routine task inputs minus the log of manual task inputs and the log of abstract task inputs. We concord this to the Census industry classification using weights based on 1980 IPUMS data. We then concord to the SIC level using a concordance from Lake and Millimet (2016). We take data on automation – specifically "displacement", or reductions in labor's share of value added after controlling for factor price movements – at the Census industry classification between 1947 and 1978 from Acemoglu and Restrepo (2019) and Acemoglu and Restrepo (2020) and concord this to the SIC level using concordances from Autor et al. (2019).

[21]As noted above, our exclusion restriction revolves around the column 2 tariffs. To this end, we note that the qualitative results from a version of Table 1 with the column 2 tariff as the dependent variable

the Swiss IV and trade-policy-related variables. The Swiss IV is positively correlated with industry-level tariff changes in trade partners, suggesting that industry-level Tokyo Round liberalization choices were correlated across countries. Further, it is negatively correlated with the MFA dummy and the importance of specific tariffs, indicating that 1978 column 2 tariffs were higher in textile industries and industries relying heavily on specific tariffs. It is also positively correlated with 1970s price growth, again underscoring the importance of specific tariffs, as 1970s price growth erodes the 1978 AVE in industries relying on specific tariffs.

Columns 2 and 4 examine the correlation between the Swiss IV and factor intensity measures. Tariff changes implied by our instrument are positively correlated with 1978 log capital-labor ratios and the non-production labor share of employment. This implies that labor-intensive and, especially, production-labor-intensive industries had higher 1978 column 2 tariffs. We also find a negative correlation with women's employment shares, consistent with industries such as textiles being female-intensive and receiving greater protection prior to the Tokyo Round. Finally, the instrument is positively correlated with the materials share of shipments, indicating that column 2 tariffs were generally lower in industries relying heavily on intermediate inputs. Column 3 of the table indicates that column 2 tariffs are negatively correlated with investment and automation, both of which proxy for the expected extent of the industry's 1980 technological change. They are positively correlated with routineness, consistent with the finding that production-labor-intensive industries faced higher column 2 tariffs. Overall, our Swiss IV, and hence the magnitude of instrumented Tokyo tariff cuts, is correlated with a host of 1978 industry characteristics.

While we will control for these characteristics when estimating equation (1), the stark and systematic correlation between these variables and our Swiss IV raises the concern that unobservable variables might also be correlated with the instrument. This is an especially important concern in the context of skill-biased technological change during the 1980s. In this spirit, we think of each 1978 industry characteristic as consisting of a component that has persisted since the Smoot-Hawley era and a component that has evolved over time. Importantly, policymakers who crafted the Smoot-Hawley tariffs are unlikely to have predicted the technological advances that led to skill-biased technological change five decades later. This suggests that identification concerns do not primarily revolve around the correlation

remain the same (although, naturally, the point estimates differ). More generally, the issue of whether the column 2 tariffs are correlated with modern industry-level characteristics is something the literature has largely ignored.

between our instrument and the components of 1978 industry characteristics driven by persistent 1930 traits. Instead, concerns relate primarily to any residual correlation between our instrument and 1978 characteristics after conditioning on 1930 traits. Such correlation would raise concerns that unobserved contemporaneous shocks, especially those related to technological change throughout the 1980s, may drive our results.

To explore this concern, we use the 1929 US Census of Manufactures to construct variables analogous to the 1978 industry characteristics included in column 4.[22] In column 5 of Table 1 we include controls for log horsepower (HP) relative to total employment, the share of employment accounted for by salaried workers (as opposed to wage earners), the share of employment accounted for by women, material inputs relative to shipments, and the log change in horsepower between 1927 and 1929. We view these, respectively, as proxies for the the log capital-labor ratio, non-production share of employment, women's share of employment, materials to shipments, and log investment. Appendix Figure A.4 documents a strong positive correlation between these variables in 1929 and 1978, reflecting the long-run persistence in these industry traits.[23] Column 5 shows that the 1929 values of materials to shipments and women's employment shares are statistically significant and of similar magnitude to the 1978 point estimates in column 4, suggesting that 1929 industry traits do play a role in explaining 1930 Smoot-Hawley tariffs and, in turn, 1978 column 2 tariffs.[24] More importantly, however, column 5 shows that none of the 1978 industry characteristics are statistically significant, and the point estimates fall substantially, once we have controlled for 1929 characteristics. This suggests that the relationship between 1978 industry traits and the Swiss IV is a product of persistence in those characteristics over time and not due to any fundamental differences between 1929 and 1978 traits.

Our identification strategy ultimately rests on the assumption that, conditional on our industry controls, unobserved determinants of 1980s import and labor market outcomes are uncorrelated with Smoot-Hawley tariffs. Our balance tests suggest this is plausible: any relationship between 1978 industry traits and our instrument is driven by persistence in 1929 characteristics, and it is unlikely policymakers set 1930 tariffs anticipating skill-biased technological change five decades later. For example, Appendix Table A.2 shows that

[22]We digitize Table 5 from Chapter II of Volume I of the 1929 Census of Manufactures. We manually concord Census industries to the SIC system.

[23]The correlation is between 0.65 and 0.75 for the log capital-labor ratio, non-production share of employment, women share of employment and materials to shipments. It is 0.21 for log investment.

[24]Consistent with Irwin and Kroszner (1996) and Irwin and Soderbery (2021), Appendix Table A.2 explicitly shows the 1930 Smoot-Hawley tariffs are correlated with contemporaneous industry characteristics.

labor-intensive and, especially, production-labor-intensive industries received greater tariff protection in 1930. Further, Table 1 shows this was still reflected in the 1978 column 2 tariffs, implying a larger liberalization under the Swiss IV for industries with higher 1978 labor-intensity and/or higher 1978 production-labor-intensity. Conditional on 1929 industry characteristics, however, 1978 column 2 tariffs are systematically unrelated to 1978 values of labor-intensity and production-labor-intensity, each a plausible proxy for 1980s skill-biased technological change.

4 Results

4.1 Trade flows

To the extent that tariffs affect labor market outcomes, they do so through their effect on goods prices and, in turn, trade flows. With this in mind, in Table 2 we explore the relationship between industry-level tariff liberalization and trade flows. Column 1 presents the effect of industry tariff changes on log import growth using an OLS specification.[25] While the point estimate has the expected sign, the magnitude is small, with an elasticity near 1.7, and the effect is not statistically significant. In column 2, we instrument for observed AVE tariff changes with our Swiss IV. Underscoring the importance of addressing endogeneity in tariff changes in this era, the results in the column change substantially. The relationship remains negative, but the elasticity rises to 9 and is significant at the 1% level. This elasticity, while large, is consistent with other analyses of trade liberalizations.[26] Further, consistent with the discussion in Section 3, the first-stage F-statistic suggests that the instrument is strong. We also note that the bias upwards toward zero evident in column 1 places structure of the nature of the endogeneity concern. Specifically, it suggests that policymakers protected industries in which they expected more rapid import growth.

Table 2 about here.

In columns 3 and 4 we examine the robustness of this result to controlling for other trade policy measures, as discussed in Table 1 above. A notable feature of the US tariff schedule in this era is the fact that many tariffs were specified in per-unit, or *specific* terms.

[25]To maintain a consistent sample size for each column, we restrict our sample to the 392 industries for which we are able to calculate all outcomes. Summary statistics for these outcomes may be found in Appendix Table A.1.

[26]See, e.g., Ruhl (2008) and Head and Mayer (2014).

Indeed, as of 1979, approximately 17% of tariff revenue in the manufacturing sector was generated by specific rather than ad valorem tariffs and roughly 70% of the industries in our sample have specific tariffs on some products. This distinction is important, as the protection afforded by specific tariffs varies inversely with the price level (Greenland et al., 2025).[27] This implies that industries may be differentially exposed to 1980s price variation as a function of dependence on specific tariffs. Additionally, in the presence of specific tariffs, 1970s price changes directly affect pre-Tokyo AVE tariffs.

To address these issues, in column 3 we control for the change in log import unit values between 1972 and 1979 and for the 1979 industry-level share of tariff-inclusive prices accounted for by specific tariffs. Inclusion of these controls increases the estimated elasticity slightly, to 9.7, but leaves the results qualitatively unchanged. In column 4 we control for more explicit trade policies of the era. First, we introduce a dummy variable equal to 1 for industries with products subject to MFA restrictions. As discussed above, textiles and apparel saw large tariff reductions under the Tokyo Round.[28] For these industries, it is important to separate tariffs from other non-tariff barriers. Second, we account for the fact that, while US import tariffs declined under the Tokyo Round, so too did the tariffs on US exports to GATT member countries. As the extent of industry-level liberalization was correlated across countries, it is possible that the estimated elasticity captures some component of the effects of tariff liberalization on US exports. We address this possibility by controlling for the extent of tariff liberalization in Canada, Japan, and the EEC. As can be seen in column 4 of the table, controlling for the MFA and tariff liberalization on US exports reduces the point estimate of interest slightly, but leaves it large and statistically significant.

It is possible that consumption responds to the Tokyo Round liberalization, and imports respond mechanically as a result. To explore this possibility, in column 5 we repeat the specification from column 4 with changes in import penetration, defined as imports relative to output consumed domestically – i.e., shipments less net exports – as our outcome of interest. We find similar results, with the Tokyo Round liberalization increasing import penetration in more exposed relative to less exposed industries. To compare magnitudes across columns 4 and 5, note than an interquartile increase in the extent of industry-level tariff liberalization reduces log import growth by approximately 37% of its interquartile range, while the analogous value for import penetration is slightly larger at 54%. Tariff

[27]For detailed discussions of the importance of specific tariffs in historical contexts, see Crucini (1994), Irwin (1998), and Greenland and Lopresti (2024).

[28]Textiles and apparel were exempted from the full Swiss-implied cuts, but still saw larger than average tariff reductions.

liberalization plays a substantial role in explaining changes in US imports, however defined, throughout the 1980s.

While our focus is the effect of the Tokyo Round on imports, it may be the case that exports respond as well, either through indirect effects of tariff liberalization on intermediate inputs or due to the multilateral nature of the agreement. To explore this possibility, columns 6 and 7 examine the impact of tariff changes on growth in log exports and exports relative to shipments, respectively. While the point estimate of interest in both columns is negative, neither is statistically significant. Further, the export elasticity in column 6 is less than half of the import elasticity in column 4. Taken together, this suggests that industry-level exports respond weakly, if at all, to Tokyo Round import tariff cuts.

Ultimately, then, Table 2 delivers two key messages. First, Tokyo Round tariff reductions in the US have large effects on imports but no statistically significant effect on exports.[29] Second, failing to account for endogenous tariff policy in the Tokyo era yields substantially biased results. In particular, policymakers appear to have protected industries that were more exposed to import growth during the liberalization, highlighting the value of our instrument in isolating random variation in tariff cuts.[30]

4.2 Within-industry pay inequality

We now turn to the impact of tariff liberalization on within-industry inequality, beginning with an analysis at the 4-digit 1987 SIC industry level using the NBER-CES Manufacturing Industry Database (Becker et al., 2021). Specifically, in Table 3 we explore how tariff changes affect industry-level log pay differences between non-production and production workers between 1979 and 1988.

Table 3 about here.

In column 1 of the table we regress changes in log pay inequality against changes in industry level $\Delta \ln(1 + AVE_i)$, instrumenting with our Swiss measure as above. The column

[29] In Appendix Table A.3 we conduct a placebo exercise by regressing trade-related outcomes from 1972-1979 on our instrument. We find that the instrumented tariff changes are not predictive of import growth in the pre-Tokyo era.

[30] There is precedent for using initial tariff levels as an instrument for the magnitude of the tariff cut – see, e.g., Goldberg and Pavcnik (2005). In Appendix Table A.4, we report estimates of import elasticities using the initial column 2 AVE tariff as an IV instead. Given the structure of the Swiss formula, these two predicted tariff changes are very highly correlated and our results are essentially unchanged by this alternative approach (this applies for our subsequent inequality results as well). This highlights the value of using the Swiss formula for understanding endogenous policymaking as we do in Appendix C.

shows that tariff liberalization leads to a relative increase in within-industry pay inequality, shifting payments towards non-production workers. In particular, an interquartile reduction in tariff levels increases the log pay differential by approximately 0.05 log points.[31]

In column 2, we introduce controls for trade policy, as discussed above. In particular, we control for changes in industry log unit values between 1972 and 1979, the share of the tariff-inclusive price accounted for by specific tariffs in 1979, whether an industry was exposed to the MFA, and the 1979-1988 change in AVE on industry exports. Inclusion of these controls increases the magnitude of the point estimate of interest considerably. The results in the column imply that an interquartile increase in industry tariff liberalization increases log pay inequality by approximately 0.09. Exposure to the MFA is particularly relevant. The point estimate on the control itself is negative and statistically significant, implying that industries covered by the MFA experienced relative reductions in log pay inequality between 1979 and 1988. Further, that its inclusion increases the point estimate on $\Delta \ln(1 + AVE_i)$ is consistent with the fact that textiles and apparel experienced large reductions in tariffs under the Tokyo Round. While the tariff liberalization itself increases pay inequality, MFA coverage reduces it. Failing to control for MFA exposure thus biases the point estimate on tariff liberalization upwards toward zero.

In column 3 we introduce controls for industry factor intensity, including log capital-to-labor ratios, the share of employment accounted for by non-production workers, the share of employment accounted for by women, and material inputs relative to shipments, all as of 1978. Including these controls leaves our result essentially unchanged and statistically significant at the 1% level. This suggests that, while Table 1 reveals that these measures are in fact correlated with our instrument, they do not separately determine inequality in the 1980s.

Finally, in column 4 we control for proxies of technological change: log industry investment as of 1978, log investment growth between 1970 and 1978, the routineness index of Autor and Dorn (2013), and a measure of automation between 1947 and 1978, following Acemoglu and Restrepo (2019, 2020). Automation and lagged investment growth, in particular, play a role in explaining rising pay inequality in our data. However, the magnitude of our point estimate of interest is unaffected: an interquartile increase in tariff liberalization increases log pay inequality by 0.08 log points.

Table 4 about here.

[31]The interquartile range in observed AVE changes and log pay inequality between 1979 and 1988 is, respectively, to 0.029 and 0.24 log points.

As discussed in Section 3, our identification strategy rests on the notion that Smoot-Hawley tariffs were as-good-as-randomly assigned conditional on observable characteristics of industries in 1929. So far we have implemented this approach by relying on 1978 column 2 tariffs, which reflect the 1930 Smoot-Hawley tariffs, for our Swiss IV and controlling for 1978 industry characteristics. Using 1978 column 2 tariffs allows us to avoid concording tariff lines that stretch over 50 years and undergo numerous changes both within the pre-1962 Schedule A classification system and the post-1963 TSUSA classification system. That said, an alternative approach is to undertake this concordance exercise and directly use the 1930 Smoot-Hawley tariffs instead of relying on the 1978 column 2 tariffs that only indirectly reflect the 1930 tariffs. An additional appeal of this approach is that, unlike 1930 AVEs, 1978 AVEs depend on 1978 prices in the presence of specific tariffs. This creates a possible endogenous relationship between our instrument and subsequent economic outcomes if those outcomes are a function of 1978 prices.

Thus, panel A of Table 4 repeats the specifications from Table 3 after constructing the Swiss IV using 1930 rather than 1978 industry-level AVEs. We use the digitized 1930 edition of *Foreign Commerce and Navigation of the United States* from Greenland and Lopresti (2024) to calculate tariff-line-level AVEs and construct changes in the Swiss IV following equations (3) and (4). We then aggregate these tariff-line measures to the SIC level. Panel A clearly shows our results are robust to this alternative specification: the implied effect of an interquartile liberalization is statistically significant and nearly double that of the baseline specification.[32]

An alternative way our exclusion restriction could be violated is if changes in 1980s pay inequality represent a continuation of pre-existing trends that are coincident with but not driven by the Tokyo Round liberalization. While the industry controls in Table 3 address this to some extent, panels B and C of Table 4 explore the possibility more directly. Panel B regresses the lagged change in pay inequality between 1972 and 1979 on the instrumented change in tariffs between 1979 and 1988, repeating the specifications from Table 3.[33] To the extent that our baseline results capture pre-existing labor market trends, we would expect similar results to Table 3. This is clearly not the case. The coefficient of interest is

[32]In Appendix Table A.5, we show the results regarding the impacts of Tokyo tariff cuts on trade outcomes in Table 2 are robust to using the instrument in panel A of Table 4 based on 1930 tariffs. In Appendix Table A.6 we show the results in Table 4 are robust to using the 1978 column 2 tariff itself as the instrument instead of the Swiss IV.

[33]We choose this period to align with our import regression, for which data are only available beginning in 1972. To facilitate comparison we scale up the 1972-1979 change to its 9-year equivalent change (i.e. multiplying by 9/7).

insignificant in columns 1 and 2 and only significant at the 10% level in column 4. Moreover, the sign of the coefficient is reversed, suggesting that inequality was *falling* throughout the 1970s in industries that would ultimately be exposed to the Tokyo Round tariff cuts.

Panel C directly accounts for pre-existing trends by re-estimating each column from Table 3 while controlling for 1972-1979 changes in log pay inequality. The results show that this has little effect on our baseline results. Our preferred estimate, in column 4, falls in magnitude from -2.9 to -2.5, but remains significant at the 5% level. This, again, suggests that our results are not driven by pre-existing trends.[34]

The results above focus on the effect of changes in own-industry tariffs, but ignore any indirect effects through upstream liberalization. To explore the role of input tariff changes in our findings, we create a weighted average of observed tariff changes on an industry's upstream inputs. An input tariff receives greater weight in the average when the industry relies more on the input for its own production, as calculated using the 1977 BEA input-output table, and the input represents a larger share of 1978 US imports. To account for endogeneity of tariff changes on inputs, we also use our Swiss IV approach to construct an instrument for upstream tariffs.[35]

Table 5 about here.

Table 5 repeats the specifications from Table 3 while additionally controlling for instrumented upstream tariffs.[36] Two points emerge from the table. First, the inclusion of upstream tariffs has little effect on our baseline estimates. In column 4, which includes the full set of baseline controls, the addition of $\Delta \ln \left(1 + AVE_{i,t}^{Up}\right)$ reduces the point estimate of interest slightly, but leaves our result qualitatively unchanged. Second, the effect of upstream tariff liberalization reinforces the effect of own-industry liberalization, albeit somewhat noisily. The estimates in column 4 imply that an interquartile reduction in upstream tariffs increases within-industry pay inequality by approximately 0.04 log points, though the effect is only statistically significant at the 10% level.

[34] In Appendix Table A.7 we show an analogous result for our trade outcomes: Tariff changes are not related to pre-Tokyo trade outcomes, and their relationship with post-Tokyo outcomes is not impacted by inclusion of lagged outcomes.

[35] We detail construction of these upstream tariff variables in Appendix B.2.

[36] We also estimate the effects of changes in upstream tariffs on trade outcomes in Appendix Table A.8.

4.3 Alternative Time Horizons and Employment v. Wage Margins

Our baseline estimates examine changes between 1979 and 1988, spanning the year immediately preceding the implementation of the Tokyo Round liberalization to the year after the phase-in had finished. In Panel A of Figure 6, we explore the relationship between the ultimate Tokyo Round tariff cut and pay inequality over different time horizons. In particular, we repeat the specification from column 4 of Table 3 separately for changes in the dependent variable between 1979 and the respective end years.

Figure 6 about here.

As is clear from panel (a) of the figure, our results are broadly robust to alternative time horizons. The estimated effects on pay inequality grow steadily through 1982 and remain stable through the remainder of the sample. This is consistent with industries adjusting over time to the phase-in of the Tokyo Round tariffs.[37]

Rising pay inequality within industries can reflect either rising wage inequality or a shift in the composition of labor towards non-production workers (i.e. rising "skill intensity"). We explore this distinction in panels (b) and (c) of Figure 6, decomposing pay inequality into its skill intensity margin, defined as the log ratio of non-production to production employment, and its wage inequality margin, defined as the log ratio of non-production labor payroll per non-production worker relative to production labor payroll per production worker. The panels reveal a dominant role for the skill intensity margin. While somewhat less precise than the estimates for pay inequality, the skill intensity estimates are negative at all time horizons save for one-year changes and are of similar magnitude to the pay inequality estimates. Point estimates on wage inequality, on the other hand, are close to zero throughout the sample and only become statistically significant at the 10% level in 1988.[38]

4.4 Inequality Within and Between Industries

Our analysis thus far emphasizes the relationship between tariff changes and pay inequality *within* industries. This, however, ignores any role for inequality *between* industries, such as might arise if exported industries expand and imported industries contract. To this end, we now use the within-between decomposition of Berman et al. (1994) to quantify the relevance

[37] Appendix Figure A.5 presents an analogous figure displaying the effects on trade outcomes at different time horizons.

[38] Appendix Table A.9 reports results analogous to Table 3 for the skill intensity and wage inequality margins.

of our within-industry analysis for observed changes in aggregate inequality.

Specifically, we decompose the observed change in the non-production share of US pay between 1979 and 1988 into its within- and between-industry components. To that end, let $y_{NP} = \frac{Pay_{NP}}{Pay}$ denote the non-production share of US manufacturing pay, $y_i = \frac{Pay_i}{Pay}$ denote industry i's share of manufacturing pay, and $y_{NP,i} = \frac{Pay_{NP,i}}{Pay_i}$ denote the non-production share of pay in industry i. Changes in aggregate pay inequality can then be decomposed as follows:

$$\Delta y_{NP} = \sum_i \bar{y}_i \Delta y_{NP,i} + \sum_i \bar{y}_{NP,i} \Delta y_i, \tag{5}$$

where $\bar{y}$ denotes the mean of variable y across the initial year 1979 and the end year 1988. The decomposition illustrates that the change in the aggregate non-production share of pay depends on two components. First, the "within component" aggregates within-industry changes in non-production pay shares, $\Delta y_{NP,i}$, using an industry's mean share of pay, $\bar{y}_i$, as the weight. Second, the "between" component aggregates changes in the between-industry shares of aggregate pay, Δy_i, using industry-specific mean non-production shares of pay, $y_{NP,i}$, as weights.

The average between-industry share of aggregate pay, $\bar{y}_i$, and the average within-industry non-production share of pay, $\bar{y}_{NP,i}$, each depend on their initial value in 1979 and the associated change between 1979 and 1988. Specifically, we can write

$$\Delta y_{NP} = \sum_i \left(y_{i,1979} + \frac{1}{2}\Delta y_{i,1979} \right) \Delta y_{NP,i} + \sum_i \left(y_{NP,i,1979} + \frac{1}{2}\Delta y_{NP,i,1979} \right) \Delta y_i. \tag{6}$$

Following equation (6), we can calculate the within and between components of the change in the aggregate non-production share of pay stemming from the liberalization. This calculation requires four pieces of information: 1979 industry pay shares ($y_{i,1979}$), 1979 within-industry non-production pay shares ($y_{NP,i,1979}$), and estimates of both the change in industry pay shares (Δy_i) and the change in within-industry non-production pay shares ($\Delta y_{NP,i}$) caused by the Tokyo Round. We can then assess the aggregate importance of the Tokyo Round for US inequality by comparing the observed change in the US non-production pay share to the change due to the Tokyo Round.

Figure 7 about here.

For the sake of space, we provide a full derivation and detail the estimation procedure in Appendix D. We report our results in Figure 7, which displays the share of aggregate changes in pay inequality accounted for by US tariff changes at time horizons between one and nine years.[39] As is clear from the figure, the Tokyo Round increases pay inequality meaningfully, accounting for between 15 and 25% of the aggregate change in pay inequality in the sample period. Further, this operates entirely through within-industry channels; the between-industry response, if anything, works to reduce pay inequality over the nine-year change, though this effect is small. Consistent with this, in Appendix Figure D.1 we show that the share of total pay accounted for by an industry does not respond to tariff liberalization.

5 Pay Inequality Heterogeneity

The results above are unambiguous: tariff liberalization throughout the 1980s led to a reallocation within industries away from production workers towards non-production workers. However, the nature of the NBER-CES manufacturing data limits our ability to parse this result further. In particular, the data do not allow us to explore heterogeneity by worker or occupation types to more fully understand the labor market consequences of the liberalization.

To address this shortcoming, we turn to Census microdata from IPUMS (Ruggles et al., 2024). Specifically, we use the 5% 1980 and 1990 samples to calculate payments among employed individuals aged 25 to 65 for each of the 72 Census manufacturing industries. We define production workers as those belonging to the following occupation groups, as defined in Autor and Dorn (2013): production and craft workers; machine operators and assemblers; and transport, construction, mechanical, mining, and farm workers. All other occupations are defined as non-production. These include clerical, administrative support and retail sales workers; managers, professional, and technical workers; and low-skill service workers.[40]

Table 6 about here.

We begin by replicating our baseline finding from column 4 of Table 3 in the IPUMS data. We report these results in column 1 of Table 6. The two samples yield similar findings: a one log point reduction in tariffs corresponds to a 3.4 log point increase in the log pay inequality

[39] As our estimation reveals relative responses across industries, we cannot explain changes in pay inequality that are common across industries.

[40] In Appendix Figure A.6 we show that the distribution of tariff changes and log pay inequality changes exhibit similar dispersion across the 4-digit 1987 SIC industries and Census industries. Summary statistics for these variables can also be found in Appendix Table A.1.

between production and non-production workers, compared to 2.9 in the NBER-CES data. In columns 2 and 3 of the table, we take advantage of the additional demographic information available in the IPUMS data to explore the effect separately for men and women. The results are striking: while the point estimate is negative for both groups, it is only statistically significant and is nearly three times as large for women.[41] That is, rising inequality in response to tariff liberalization seems to be driven overwhelmingly by women.[42]

These results suggest important differences in the nature of employment for men and women in this era. Indeed, Appendix Figure A.8 and Appendix Table A.13 show that women's manufacturing employment is substantially more concentrated than men's and, unlike men, is heavily concentrated in routine occupations. Women's manufacturing employment is highly skewed toward occupations such as textile sewing machine operators, equipment assemblers, and secretaries. More generally, the seven occupations that employ the most women account for more than 40% of women's employment and six of them have above average routineness. Conversely, men are much more likely to work as production supervisors, machine operators, and managers. Five of the six occupations that employ the most men have below average routineness.[43]

Given the heterogeneity in the importance of routine occupations between men and women, in panel A of Table 7 we repeat the specifications from Table 6 separately for occupations above and below the median routineness index level. Here, too, we observe a striking pattern: among more routine occupations, there is a shift in pay towards non-production workers among both men and women. While the effect for women is larger than that for men, both are large and statistically significant. Conversely, neither group exhibits a statistically significant response among non-routine occupations. That is, the effects of tariff liberalization on inequality are driven exclusively by routine occupations, with an effect that falls on both men and women, albeit more heavily on women. This suggests a possible

[41] In Appendix Figure A.7 we display the kernel density of changes in log pay inequality for men and women separately. Consistent with the summary statistics in Appendix Table A.1, this distribution has a higher mean and variance for women than for men.

[42] In Appendix Table A.10 we present estimates in which we sequentially introduce controls as in Table 3. The differential effect among women is apparent in all specifications. Additionally, in Appendix Table A.11 we conduct "placebo" exercises and control for lagged outcomes as in Table 4. Finally, in Appendix Table A.12 we report similar results exploring changes in log payments to college graduates relative to non-college graduates. Here, too, we find that the increase in inequality among women is substantially larger than the average effect, though the estimated effect for women has a p-value of 0.101.

[43] At the occupation group level, approximately 80% of women's employment is accounted for by the three most routine occupation groups: machine operators, craft producers, and clerical workers. Conversely, nearly 50% of men's employment is accounted for by the three least routine occupation groups: transportation workers, managers, and low-skill service workers.

explanation for the stark heterogeneity in Table 6: women are disproportionately employed in routine occupations in the era, and it is these occupations that experience the greatest rise in pay inequality in response to the Tokyo Round.

Table 7 about here.

In a similar spirit, we explore heterogeneity along another dimension: exposure to automation. Technological change is commonly cited as a potential driver of inequality in this era, and indeed its explanatory power is often directly contrasted with that of trade. The results above suggest that trade and technology may have jointly played a role in determining the shape of inequality in the era. In particular, it is possible that industries differentially exposed to technological change responded differently to trade liberalization. We explore this possibility more directly in panel B of Table 7, in which we repeat the specifications from panel A separately for industries above and below the median level of automation (Acemoglu and Restrepo, 2019, 2020).

Again, we observe a clear pattern: among industries at or above the median level of automation between 1947 and 1978, the Tokyo Round liberalization dramatically increases pay inequality. Specifically, an interquartile increase in the liberalization raises pay inequality by 0.16 log points, with an effect that is statistically significant for both men and women. Conversely, among industries below the median level of automation we observe a statistically significant effect of the liberalization only among women, for whom the estimated effect is 40% lower than in high automation industries. This suggests that, rather than operating in isolation, tariff liberalization and technological change interact to drive inequality in this era.

These findings underscore a broader point: the effect of the Tokyo Round on inequality was not uniform, and was shaped jointly by both trade and technological change. Workers and industries exposed to both falling tariffs and shifting technology experienced larger increases in the relative pay of non-production workers, while those only exposed to tariff liberalization saw more muted effects. It was the combination of trade and technology, rather than either in isolation, that reshaped labor markets in the era.

6 Conclusion

Despite representing a tariff liberalization larger in magnitude than any of CUSFTA, NAFTA, or granting PNTR to China, the Tokyo Round of the GATT has received relatively little empirical attention. In this paper, we exploit newly available data and a novel identification strategy to explore the trade and distributional consequences of the Tokyo Round. We show that the formula-based nature of Tokyo Round tariff negotiations, combined with the historical importance of 1930s tariffs within the institutional structure of the US tariff schedule, allows construction of a simple IV that circumvents standard concerns about endogenous tariff protection.

We use our resulting "Swiss IV" to show that Tokyo Round tariff cuts led to rising 1980s import growth and were associated with an increase in pay inequality between production and non-production workers that is overwhelmingly caused by changes in skill intensity, rather than rising wage inequality. This effect is driven by routine occupations and industries most exposed to technological change. Strikingly, the increases in inequality fall disproportionately on women, who are over-represented in routine manufacturing jobs in this era.

The Tokyo Round represents an interesting laboratory for future work given the magnitude of the liberalization alongside the clean and transparent identification strategy provided by our Swiss IV. Indeed, this is an era of falling manufacturing employment, rising women's labor force participation, declining unionization, and enormous shifts between workers of differing education levels. Each of these trends is potentially related, either directly or indirectly, to the Tokyo Round tariff liberalization.

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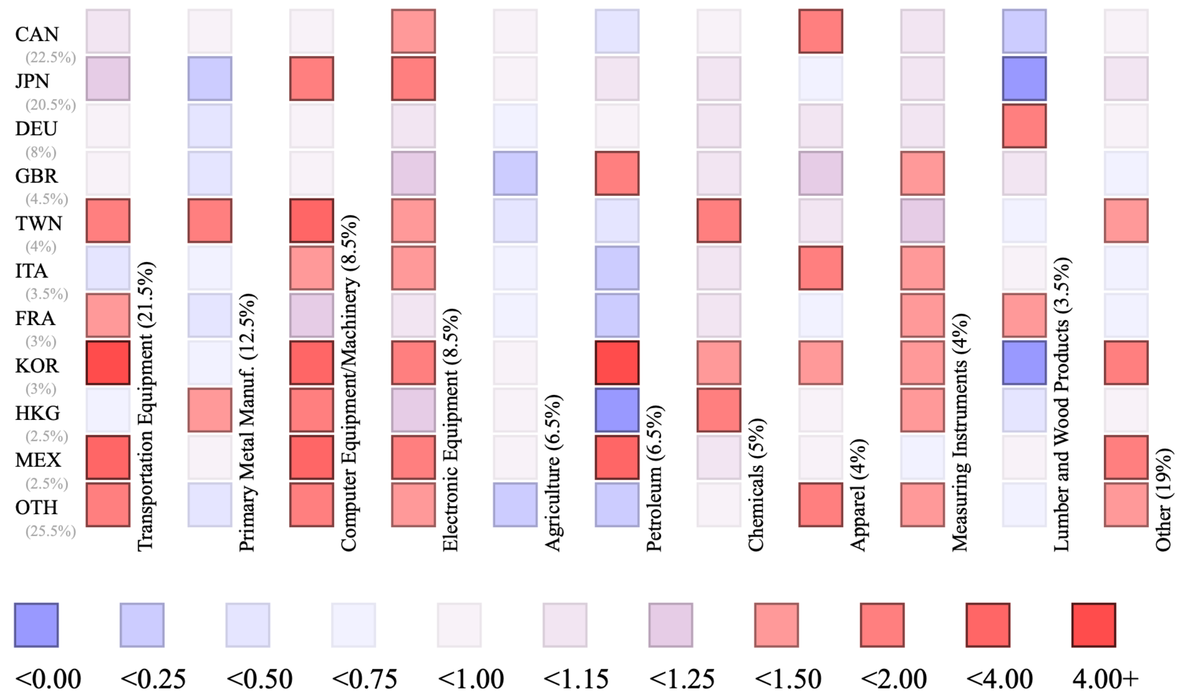
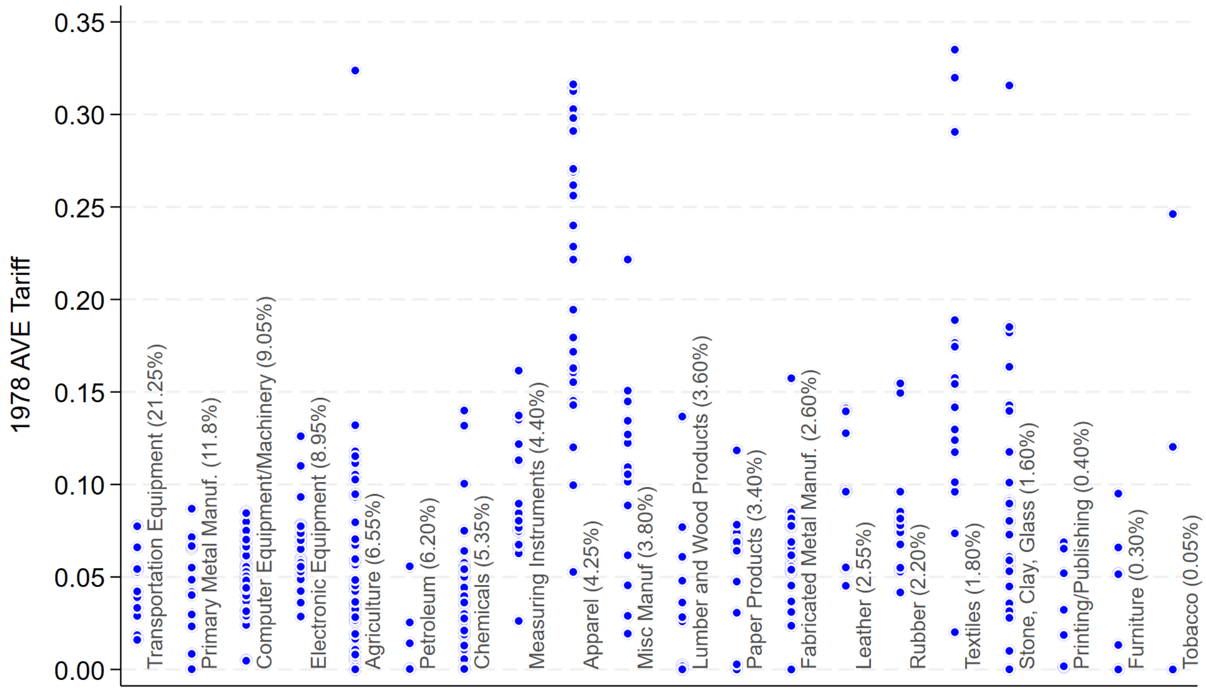
Tables and Figures

Figure 1: Import Growth by Country and Sector



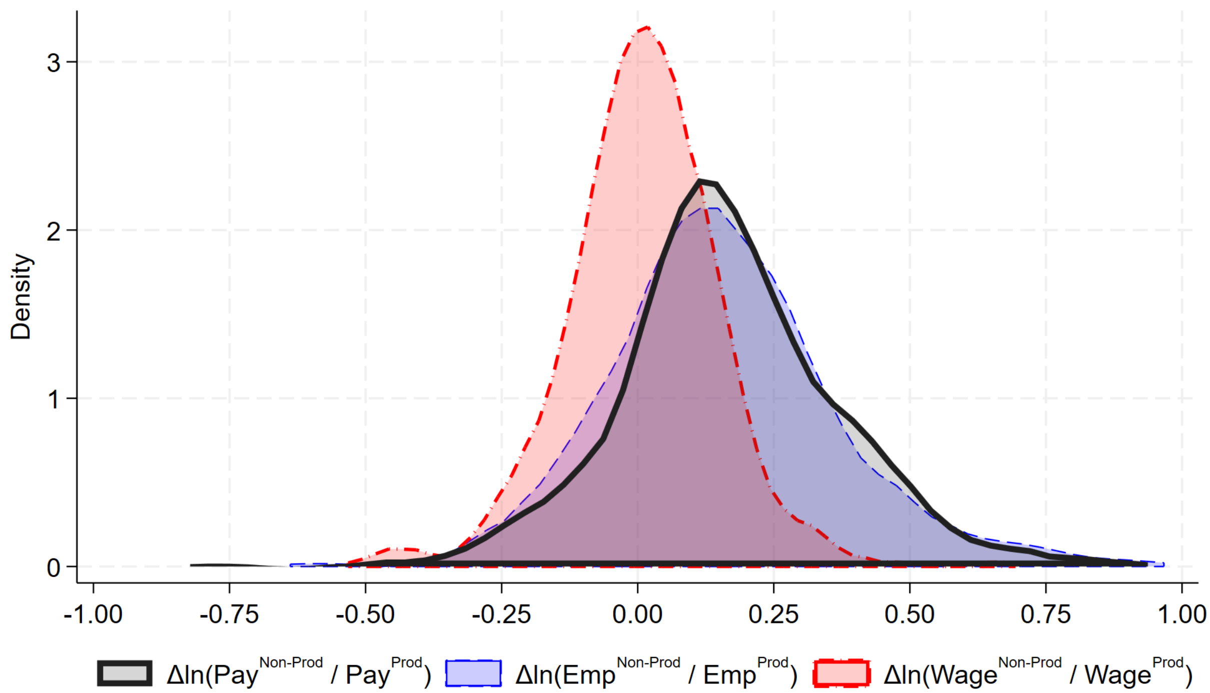
Notes: Each cell represents US log import growth between 1979 and 1988 from an exporting country in a 2-digit 1987 SIC sector. The numbers in parentheses on the vertical and horizontal axes are, respectively, shares of total US imports in 1978 accounted for by the exporter and 2-digit SIC sector. Import data taken from Schott (2008).

Figure 2: Pre-Tokyo 1978 AVE Tariffs by 2-digit SIC Sector



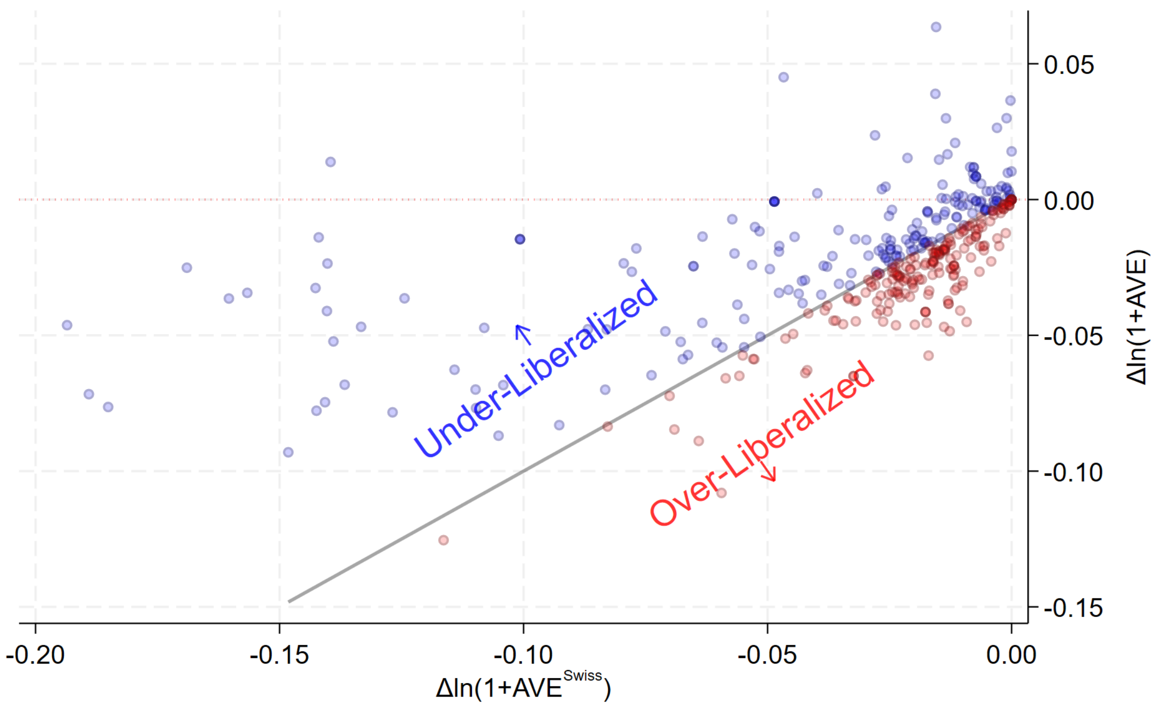
Notes: Each dot represents the 1978 AVE tariff of a 4-digit 1987 SIC industry. Dots clustered by two-digit SIC sectors and ordered by 1978 share of US imports, which are indicated in parentheses. Sectors are ordered by 1978 import share. Tariff data from Greenland et al. (2025). Import data from Schott (2008).

Figure 3: Variation in Margins of Industry Pay Inequality



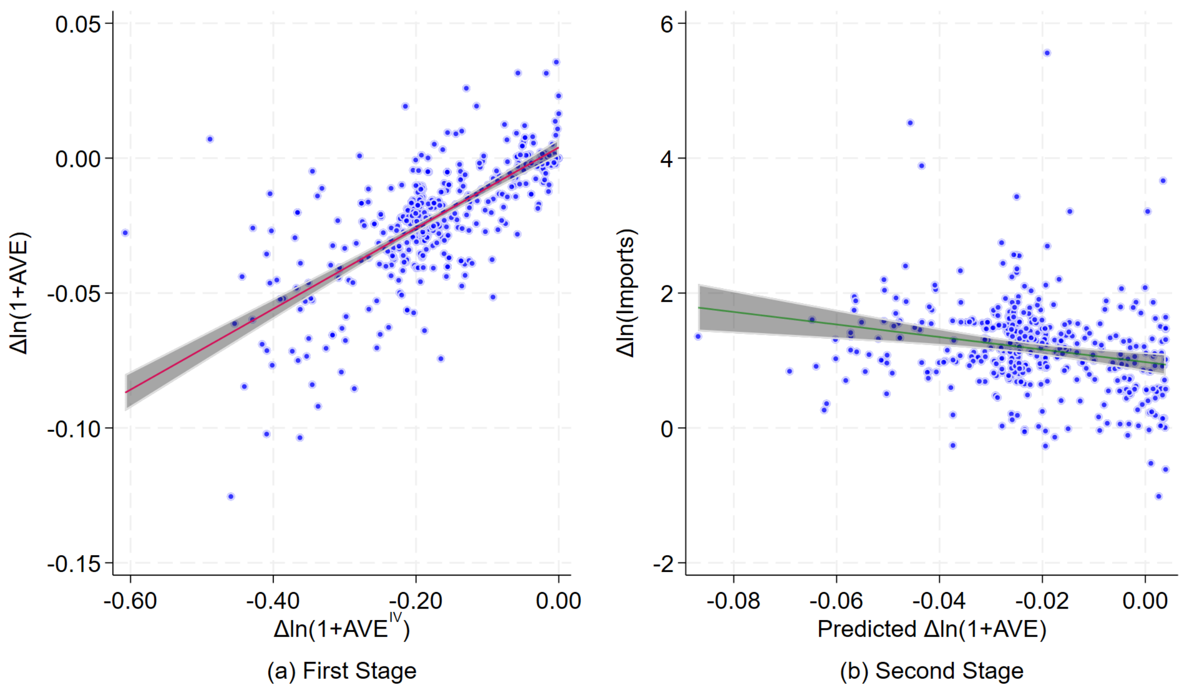
Notes: Figure displays a kernel density of the industry-level change in log pay to non-production workers relative to production workers between 1979 and 1988 in thick black. Blue (with a thin dash) and red (with a thick dash) densities decompose this log change into log changes in relative employment and relative wage margins respectively. Data from the NBER-CES Manufacturing Industry Database (Becker et al. (2021)).

Figure 4: AVE Tariff Cuts During the Tokyo Round



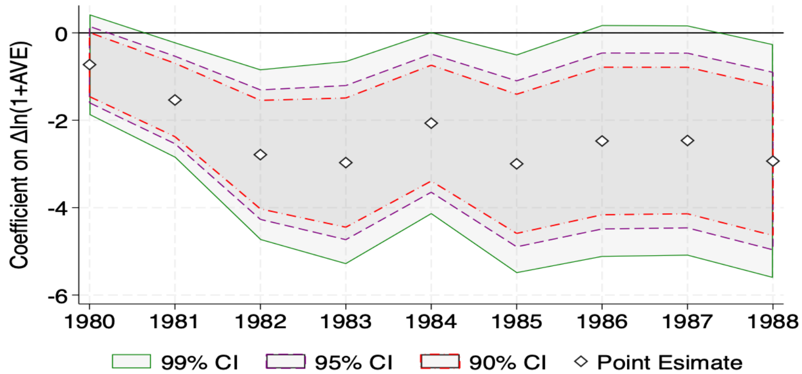
Notes: Figure displays a scatterplot at the 4-digit 1987 SIC industry-level of observed tariff changes $\Delta\ln(1+AVE)$ from 1979-1988 versus the tariff changes implied by the Swiss formula $\Delta\ln\left(1+AVE^{Swiss}\right)$. The 45 degree line in gray indicates strict adherence to the Swiss formula. Points above this line (indicated in blue) received a smaller tariff cut than prescribed while those below the line (indicated in red) received a larger tariff cut than prescribed. Tariff data from Greenland et al. (2025).

Figure 5: Effects of Tokyo Tariff Cuts on Import Growth Using Swiss IV

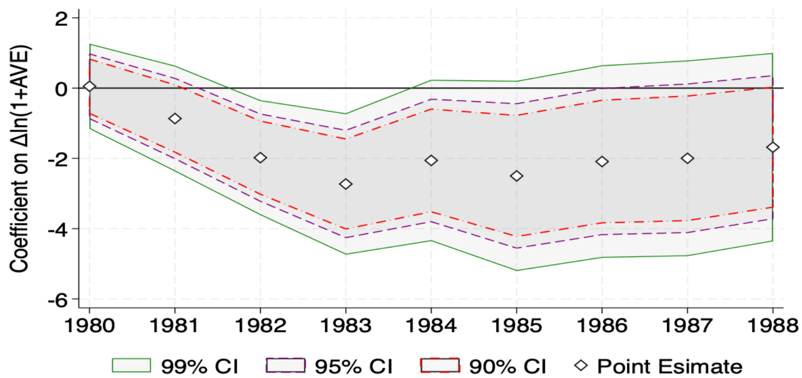


Notes: Scatterplot displaying two-stage least squares regression at 4-digit SIC industry level of import growth on changes in observed AVE tariffs $\Delta \ln(1 + AVE)$ instrumented by our Swiss IV $\Delta \ln(1 + AVE^{IV})$ as defined in equations (3) and (4). Long differences taken from 1979-1988. Two points omitted from this figure for improved readability. Results are not dependent on their omission. Tariff data from Greenland et al. (2025). Import data from Schott (2008).

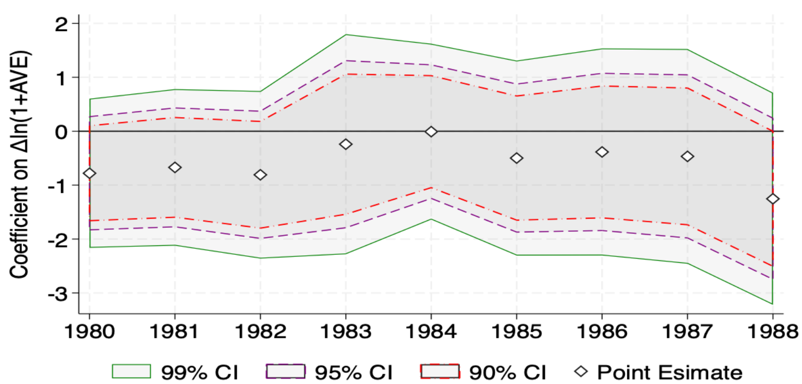
Figure 6: Tariff Liberalization and Pay Inequality – Alternative Time Horizons and Decomposition



(a) $\Delta \ln \left(\frac{Pay^{Non-Prod}}{Pay^{Prod}} \right)$



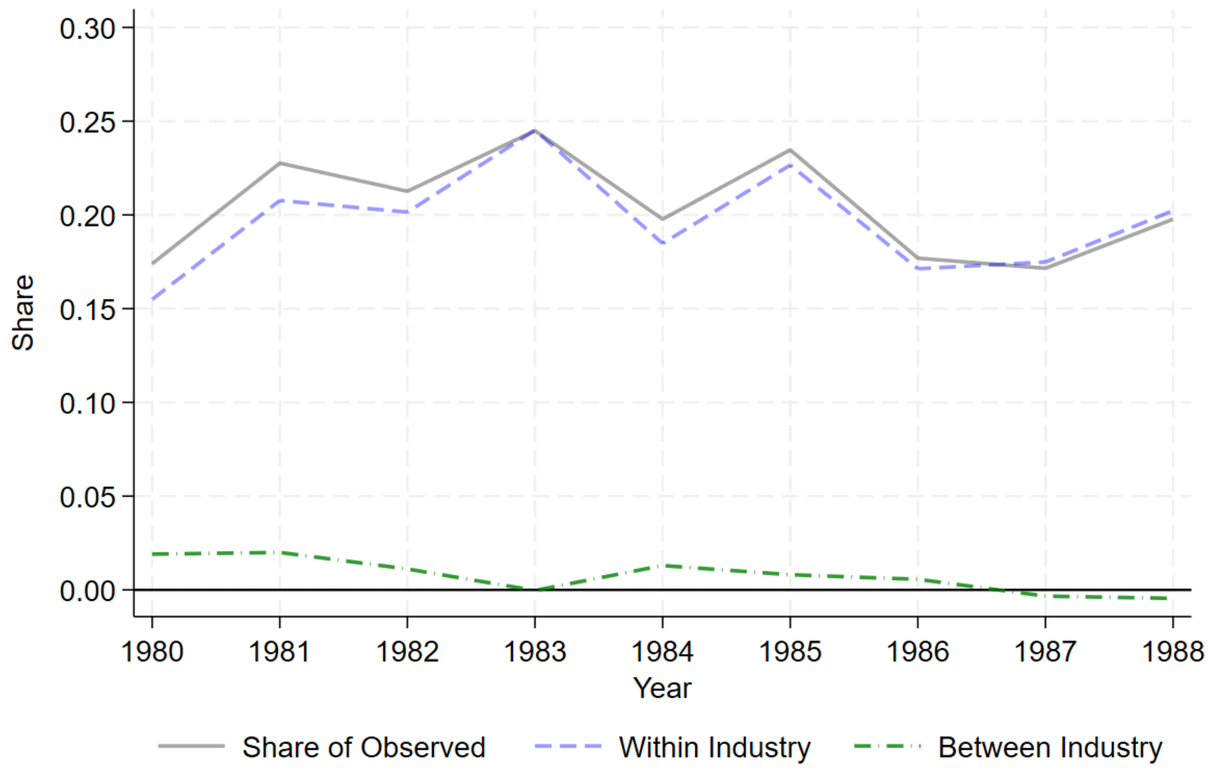
(b) $\Delta \ln \left(\frac{Emp^{Non-Prod}}{Emp^{Prod}} \right)$



(c) $\Delta \ln \left(\frac{Wage^{Non-Prod}}{Wage^{Prod}} \right)$

Notes: Figure presents estimates from estimating equation (1). Each year t point estimate is from a separate specification with the change in the dependent variable defined as the change between 1979 and year t. Dependent variable in panel (a) is change in log pay inequality between non-production and production workers. Panels (b) and (c) decompose this effect into relative employment and relative wage margins. Data measured at or concorded to the 4-digit SIC87 industry classification. Specifications use full controls as in column 4 of Table 3. Robust standard errors. Data from Becker et al. (2021) and Greenland et al. (2025).

Figure 7: Aggregate Effects of Tokyo Tariff Cuts on Within and Between Industry Pay Inequality



Notes: Figure presents estimates of the share of observed aggregate pay inequality explained by Tokyo Round tariff liberalization (solid gray line) and decomposes this share into within industry (dashed blue line) and between industry (dashed green line) components. Methodology follows Berman et al. (1994). See Appendix D for details on implementation and estimation.

Table 1: Correlates of Swiss IV

	$\Delta \ln(1 + AVE_i^{IV})$				
	(1)	(2)	(3)	(4)	(5)
$\Delta \ln(p^*_{i,t-1})$	0.089***			0.036**	0.039**
	(0.020)			(0.018)	(0.016)
$I(MFA_i)$	-0.090***			-0.037**	-0.035**
	(0.014)			(0.015)	(0.015)
$STS_i * \frac{AVE_i}{1+AVE_i}$	-2.100***			-2.048***	-1.876***
	(0.455)			(0.413)	(0.364)
$\Delta AVE_i{}^{Exports}$	0.093***			0.052*	0.041*
	(0.033)			(0.028)	(0.022)
$\frac{Capital_i}{Labor_i}$		0.013		0.023***	0.013
		(0.009)		(0.008)	(0.010)
$\frac{Emp_i^{Non-Prod}}{Emp_i}$		0.178***		0.092*	0.067
		(0.050)		(0.048)	(0.053)
$\frac{Emp_i^{Women}}{Emp_i}$		-0.249***		-0.081*	0.006
		(0.039)		(0.043)	(0.045)
$\frac{Materials_i}{Shipments_i}$		0.226***		0.203***	0.056
		(0.048)		(0.045)	(0.050)
$\ln(Investment_i)$			0.023***	0.003	0.006
			(0.005)	(0.004)	(0.004)
$\Delta \ln(Investment_{i,t-1})$			0.014	0.010	0.011
			(0.010)	(0.008)	(0.007)
$Routineness_i$			-0.100***	-0.023	-0.018
			(0.025)	(0.022)	(0.020)
$Automation_i$			0.150*	0.029	0.040
			(0.078)	(0.055)	(0.062)
$\ln(\frac{HP_{i,1929}}{Employment_{i,1929}})$					0.008
					(0.009)
$\frac{Emp_{i,1929}^{Salaried}}{Emp_{i,1929}}$					0.105
					(0.092)
$\frac{Emp_{i,1929}^{Women}}{Emp_{i,1929}}$					-0.131***
					(0.046)
$\frac{Materials_{i,1929}}{Shipments_{i,1929}}$					0.194***
					(0.047)
$\Delta \ln(HP_{i,1927-1929})$					-0.027
					(0.030)
Constant	-0.113***	-0.270***	-0.026	-0.265***	-0.286***
	(0.011)	(0.037)	(0.035)	(0.050)	(0.051)
Observations	334	334	334	334	334
R^2	0.332	0.381	0.180	0.501	0.566

Notes: Table presents correlation between industry characteristics and our instrument $\Delta \ln(1 + AVE_i^{IV})$ defined in equations (3) and (4). By definition $\Delta \ln(1 + AVE_i^{IV}) \leq 0$. Changes are contemporaneous to the outcome unless indicated otherwise. Changes subscripted $t-1$ indicate a change from 1972-1979. All data measured at or concorded to the 4-digit 1987 SIC system. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table 2: Tariff Liberalization and Trade

	(1) $\Delta\ln(Imp_i)$	(2) $\Delta\ln(Imp_i)$	(3) $\Delta\ln(Imp_i)$	(4) $\Delta\ln(Imp_i)$	(5) $\Delta ImpPen_i$	(6) $\Delta\ln(Exp_i)$	(7) $\Delta\frac{Exp_i}{Ship_i}$
$\Delta\ln(1+AVE_i)$	-1.683	-8.968***	-9.661***	-8.788***	-1.562***	-3.548	-0.333
	(3.480)	(2.573)	(2.902)	(3.203)	(0.335)	(2.173)	(0.277)
$\Delta\ln(p^*_{i,t-1})$			-0.255**	-0.249**	0.003	-0.154	-0.021
			(0.126)	(0.127)	(0.017)	(0.124)	(0.025)
$STS_i * \frac{AVE_i}{1+AVE_i}$			-2.023	-1.271	-0.846***	1.671	-0.076
			(1.779)	(1.961)	(0.198)	(1.450)	(0.096)
$I(MFA_i)$				-0.045	0.017	-0.090	-0.012*
				(0.085)	(0.015)	(0.074)	(0.007)
$\Delta AVE_i{}^{Exports}$				-0.264	0.025	0.075	0.003
				(0.251)	(0.017)	(0.091)	(0.010)
Constant	1.122***	0.963***	0.979***	0.927***	0.043***	0.446***	0.010
	(0.105)	(0.083)	(0.085)	(0.107)	(0.008)	(0.051)	(0.006)
IQE	-0.071	-0.263	-0.283	-0.258	-0.046	-0.104	-0.010
Estimator	OLS	IV	IV	IV	IV	IV	IV
Obs.	392	392	392	392	392	392	392
1^{st} Stage Coeff.		0.149	0.150	0.143	0.143	0.143	0.143
KP F-Stat	.	177.841	153.342	115.979	115.979	115.979	115.979

Notes: Table presents estimates of equation (1) where the dependent variable is the change in log imports in columns 1-4, change in import penetration, which is imports divided by shipments minus exports plus imports, in column 5, change in log exports in column 6, and change in exports scaled by shipments in column 7. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented in columns 2-7 with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data from Schott (2008) and the NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta\ln(1+AVE)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table 3: Tariff Liberalization and Pay Inequality

	$\Delta\ln(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}})$			
	(1)	(2)	(3)	(4)
$\Delta\ln(1+AVE_i)$	-1.804***	-2.994***	-2.839***	-2.934***
	(0.691)	(0.824)	(1.012)	(1.014)
$\Delta\ln(p^*_{i,t-1})$		-0.038	-0.038	-0.039
		(0.046)	(0.047)	(0.045)
$STS_i * \frac{AVE_i}{1+AVE_i}$		-0.605	-0.423	-0.219
		(0.572)	(0.624)	(0.623)
$I(MFA_i)$		-0.076***	-0.085***	-0.075**
		(0.029)	(0.030)	(0.031)
$\Delta AVE_i^{Exports}$		0.033	0.033	0.046
		(0.037)	(0.036)	(0.034)
$\frac{Capital_i}{Labor_i}$			-0.023	-0.001
			(0.020)	(0.022)
$\frac{Emp_i^{Non-Prod}}{Emp_i}$			0.022	-0.005
			(0.113)	(0.113)
$\frac{Emp_i^{Women}}{Emp_i}$			-0.030	0.050
			(0.100)	(0.109)
$\frac{Materials_i}{Shipments_i}$			0.070	0.071
			(0.115)	(0.115)
$\ln(Investment_i)$				-0.009
				(0.012)
$\Delta\ln(Investment_{i,t-1})$				0.042*
				(0.023)
$Routineness_i$				0.024
				(0.046)
$Automation_i$				0.289*
				(0.164)
Constant	0.127***	0.133***	0.194*	-0.009
	(0.019)	(0.021)	(0.102)	(0.140)
IQE	-0.053	-0.088	-0.084	-0.086
Estimator	IV	IV	IV	IV
Controls	None	+Trade	+Prod.	+Tech.
Obs.	395	395	395	395
1^{st} Stage Coeff.	0.150	0.144	0.135	0.135
KP F-Stat	180.332	119.386	93.062	91.021

Notes: Table presents estimates of equation (1) where the dependent variable is the change in log pay to non-production workers relative to production workers from 1979-1988. Column 1 has no controls. Column 2 adds the trade controls from column 1 of Table 1. Column 3 adds the production controls from column 2 of Table 1. Column 4 adds the technology controls from column 3 of Table 1. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data from the NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta\ln(1+AVE)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table 4: Tariff Liberalization and Pay Inequality – Robustness

	(1)	(2)	(3)	(4)
	Panel A: $\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}} \right)_t$			
$\Delta \ln (1 + AVE_i)$	-2.585**	-4.450***	-5.194**	-5.256**
	(1.023)	(1.638)	(2.296)	(2.314)
IQE	-0.072	-0.124	-0.145	-0.147
Obs.	312	312	312	312
1^{st} Stage Coeff.	0.104	0.072	0.057	0.058
KP F-Stat	115.767	46.561	27.096	27.687
	Panel B: $\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}} \right)_{t-1}$			
$\Delta \ln (1 + AVE_i)$	0.187	1.238	1.921**	1.827*
	(0.522)	(0.620)	(0.741)	(0.736)
IQE	0.004	0.028	0.044	0.042
Obs.	394	394	394	394
1^{st} Stage Coeff.	0.150	0.144	0.135	0.136
KP F-Stat	180.335	119.304	93.063	89.762
	Panel C: $\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}} \right)_t$			
$\Delta \ln (1 + AVE_i)$	-1.764***	-2.756***	-2.446**	-2.458**
	(0.682)	(0.811)	(1.010)	(0.992)
$\Delta \ln (\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}})_{t-1}$	-0.212**	-0.197**	-0.204**	-0.207**
	(0.124)	(0.123)	(0.124)	(0.123)
IQE	-0.052	-0.081	-0.072	-0.072
Obs.	394	394	394	394
1^{st} Stage Coeff.	0.150	0.144	0.135	0.136
KP F-Stat	177.851	115.526	88.769	86.250
Estimator	IV	IV	IV	IV
Controls	None	+Trade	+Prod.	+Tech.

Notes: Table presents estimates of equation (1) where the dependent variable is the change in log pay to non-production workers relative to production workers. Column 1 has no controls. Column 2 adds the trade controls from column 2 of Table 3. Column 3 adds the production controls from column 3 of Table 3. Column 4 adds the technology controls from column 4 of Table 3. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data from the NBER-CES Manufacturing Industry Database (Becker et al., 2021). Panel A uses 1930 AVE tariffs instead of 1978 column 2 tariffs in construction of the instrument. Panel B uses log pay inequality measured over the 1972-1979 period as the dependent variable. Panel C uses log pay inequality measured over the 1972-1979 period as an additional control. All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta \ln(1 + AVE)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table 5: Tariff Liberalization and Pay Inequality – Upstream Tariffs

	$\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}} \right)$			
	(1)	(2)	(3)	(4)
$\Delta \ln (1 + AVE_i)$	-1.628**	-2.426***	-2.574**	-2.503**
	(0.765)	(0.872)	(1.006)	(1.000)
$\Delta \ln \left(1 + AVE_i^{Up}\right)$	-1.138	-5.222*	-7.432**	-7.610*
	(2.893)	(2.957)	(3.784)	(3.952)
IQE	-0.048	-0.071	-0.076	-0.074
IQE^{up}	-0.006	-0.030	-0.042	-0.043
Obs.	394	394	394	394
Under ID p-stat	0.000	0.000	0.000	0.000
KP F-Stat	76.007	55.748	45.244	42.843
Estimator	IV	IV	IV	IV
Controls	None	+Trade	+Prod.	+Tech.

Notes: Table presents estimates of equation (1) where the dependent variable is the change in log pay to non-production workers relative to production workers from 1979-1988. Column 1 has no controls. Column 2 adds the trade controls from column 2 of Table 3. Column 3 adds the production controls from column 3 of Table 3. Column 4 adds the technology controls from column 4 of Table 3. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented with the Swiss formula implied tariff change as in equations (3)-(4) and Appendix B.2. Outcome data from the NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE and IQE^{up} indicate the effect on the dependent variable of an interquartile increase in $\Delta \ln(1 + AVE)$ and $\Delta \ln(1 + AVE^{up})$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table 6: Tariff Liberalization and Pay Inequality – Heterogeneity by Gender

	$\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}} \right)$		
	(1)	(2)	(3)
	All	Men	Women
$\Delta \ln (1 + AVE_i)$	-3.373**	-2.661	-6.457***
	(1.705)	(1.967)	(2.174)
IQE	-0.093	-0.074	-0.178
Estimator	IV	IV	IV
Obs.	72	72	72
1^{st} Stage Coeff.	0.131	0.131	0.131
KP F-Stat	33.857	33.857	33.857

Notes: Table presents estimates of equation (1) where the dependent variable is the change in log pay to non-production workers relative to production workers from 1980-1990. Column 1 uses the full sample of workers. Columns 2 and 3 restrict the sample to male and female workers respectively. All specifications use all controls from Table 3. Data measured at or concorded to the Census industry classification. AVE tariffs from Greenland et al. (2025) and are instrumented with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data from IPUMS (Ruggles et al., 2024). First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta \ln(1 + AVE_i)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table 7: Tariff Liberalization and Pay Inequality – Heterogeneity by Routineness and Automation

	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: Routine vs Non-Routine Occupations					
	Routine			Non-Routine		
	All	Men	Women	All	Men	Women
$\Delta\ln(1+AVE_i)$	-4.618**	-3.868*	-7.809***	-1.688	-1.649	-0.928
	(1.980)	(2.175)	(2.311)	(1.981)	(2.224)	(3.273)
IQE	-0.128	-0.107	-0.216	-0.047	-0.046	-0.026
Estimator	IV	IV	IV	IV	IV	IV
Obs.	72	72	72	72	72	72
1^{st} Stage Coeff.	0.131	0.131	0.131	0.131	0.131	0.131
KP F-Stat	33.857	33.857	33.857	33.857	33.857	33.857
	Panel B: High vs Low Automation Industries					
	High Automation			Low Automation		
	All	Men	Women	All	Men	Women
$\Delta\ln(1+AVE_i)$	-5.927***	-5.519**	-6.059**	-1.155	-0.895	-3.646*
	(2.160)	(2.740)	(2.762)	(1.733)	(2.031)	(1.986)
IQE	-0.164	-0.152	-0.167	-0.032	-0.025	-0.101
Estimator	IV	IV	IV	IV	IV	IV
Obs.	38	38	38	34	34	34
1^{st} Stage Coeff.	0.166	0.166	0.166	0.134	0.134	0.134
KP F-Stat	13.609	13.609	13.609	7.533	7.533	7.533

Notes: Table presents estimates of equation (1) where the dependent variable is the change in log pay to non-production workers relative to production workers from 1980-1990 among all workers (columns 1 and 4), men (column 2 and 5), and women (columns 3 and 6). Panel A presents estimates among occupations with above median routineness (i.e. routine) in columns 1-3 and below median routineness occupations (i.e. non-routine) in columns 4-6. Panel B presents estimates among industries with above median automation in columns 1-3 and below median automation in columns 4-6. All specifications use all controls from Table 3. Data measured at or concorded to the Census industry classification. AVE tariffs from Greenland et al. (2025) and are instrumented with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data from IPUMS (Ruggles et al., 2024). Routineness data from Autor and Dorn (2013). Automation data from Acemoglu and Restrepo (2019, 2020). First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta\ln(1+AVE_i)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Appendix

A Appendix Tables and Figures

Table A.1: Summary Statistics

	Obs	Mean	SD	IQR
	Panel A. SIC Analysis			
$\Delta\ln(1+AVE_i^{IV})$	392	-0.174	0.109	0.135
$\Delta\ln(1+AVE_i)$	392	-0.022	0.023	0.029
$\Delta\frac{Exports_{i,t}}{Shipments_{i,t}}$	392	0.012	0.061	0.027
ΔImport Penetration$_i$	392	0.066	0.086	0.085
$\Delta\ln(Exports_i)$	392	0.491	0.507	0.634
$\Delta\ln(Imports_i)$	392	1.159	0.837	0.694
$\Delta\ln(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}})$	392	0.165	0.213	0.239
$\Delta\ln(\frac{Emp_i^{Non-Prod}}{Emp_i^{Prod}})$	392	0.155	0.209	0.246
$\Delta\ln(\frac{Wage_i^{Non-Prod}}{Wage_i^{Prod}})$	392	0.011	0.139	0.162
	Panel B. Census Analysis			
$\Delta\ln(1+AVE_i^{IV})$	72	-0.153	0.094	0.140
$\Delta\ln(1+AVE_i)$	72	-0.020	0.020	0.028
$\Delta\ln(\frac{Pay_i^{Non-Prod,A}}{Pay_i^{Prod,A}})$	72	0.231	0.147	0.174
$\Delta\ln(\frac{Pay_i^{Non-Prod,M}}{Pay_i^{Prod,M}})$	72	0.193	0.157	0.181
$\Delta\ln(\frac{Pay_i^{Non-Prod,W}}{Pay_i^{Prod,W}})$	72	0.288	0.254	0.332

Notes: Panel A summary statistics aggregated to or recorded at the 4-digit 1987 SIC industry. Panel B summary statistics aggregated or concorded to the Census mnuafacturing industries in IPUMS (Ruggles et al., 2024). Trade flow data from Schott (2008). Tariff-line level ad valorem equivalent (AVE) tariff data from Greenland et al. (2025). Remaining SIC analysis data from the NBER-CES Manufacturing Industry Database (Becker et al., 2021) and remaining census analysis data from IPUMS (Ruggles et al., 2024). $\Delta\ln(1+AVE_i^{IV})$ is the Swiss IV defined by equations (3) and (4). Superscripts A, M, W denote the variable relates to all workers, men workers, and women workers.

Table A.2: Balance Tests and 1930 AVE Tariffs

	$\Delta_i^{1930})$					
	(1)	(2)	(3)	(4)	(5)	(6)
$\ln(\frac{HP_{i,1929}}{Employment_{i,1929}})$	-0.056***					-0.018*
	(0.006)					(0.010)
$\frac{Emp_{Salaried,i,1929}}{Emp_{i,1929}}$		-0.526***				-0.467***
		(0.105)				(0.116)
$\frac{Emp_{Women,i,1929}}{Emp_{i,1929}}$			0.352***			0.256***
			(0.037)			(0.054)
$\frac{Materials_{i,1929}}{Shipments_{i,1929}}$				-0.169***		-0.152***
				(0.055)		(0.053)
$\Delta\ln(HP_{i,1927-1929})$					-0.027	0.046
					(0.048)	(0.043)
Constant	0.347***	0.383***	0.218***	0.379***	0.304***	0.393***
	(0.009)	(0.020)	(0.012)	(0.028)	(0.012)	(0.039)
Observations	277	277	277	277	277	277
R^2	0.186	0.078	0.213	0.036	0.001	0.312

Notes: Table presents correlation between 1929 industry characteristics and our instrument $\Delta\ln(1+AVE_i^{IV})$ defined in equations (3) and (4) except that $AVE_{g,1930}$ replaces $AVE_{g,1978}^{Col2}$. Control variable data from Table 5 of Chapter II of Volume I of the 1929 Census of Manufactures. 1930 tariff data from Greenland and Lopresti (2024). All data concorded to the 4-digit 1987 SIC system. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.3: Tariff Liberalization and Trade – 1972-1979 Placebo Period

	(1) $\Delta \ln(Imp_i)$	(2) $\Delta \ln(Imp_i)$	(3) $\Delta \ln(Imp_i)$	(4) $\Delta \ln(Imp_i)$	(5) $\Delta ImpPen_i$	(6) $\Delta \ln(Exp_i)$	(7) $\Delta \frac{Exp_i}{Ship_i}$
$\Delta \ln(1 + AVE_i)$	2.307	4.934	2.785	0.525	-0.289	-2.772	0.197
	(2.040)	(3.462)	(3.339)	(3.309)	(0.383)	(3.155)	(0.414)
$\Delta \ln(p^*_{i,t-1})$			-1.038***	-0.989***	-0.001	0.357	0.001
			(0.350)	(0.356)	(0.013)	(0.230)	(0.020)
$STS_i * \frac{AVE_i}{1+AVE_i}$			-6.656**	-7.189***	-0.350**	-2.585	-0.226
			(2.727)	(2.756)	(0.173)	(2.236)	(0.165)
$I(MFA_i)$				-0.094	0.027**	0.169*	0.000
				(0.132)	(0.013)	(0.100)	(0.007)
$\Delta AVE_i{}^{Exports}$				0.236	0.024*	0.207	-0.000
				(0.187)	(0.014)	(0.170)	(0.009)
Constant	1.541***	1.599***	1.663***	1.700***	0.032***	1.690***	0.045***
	(0.064)	(0.083)	(0.087)	(0.097)	(0.010)	(0.090)	(0.012)
IQE	0.098	0.145	0.082	0.015	-0.008	-0.081	0.006
Estimator	OLS	IV	IV	IV	IV	IV	IV
Obs.	389	389	389	389	392	390	392
1^{st} Stage Coeff.		0.149	0.150	0.143	0.143	0.143	0.143
KP F-Stat	.	176.813	151.711	113.723	115.979	111.301	115.979

Notes: Table presents estimates of equation (1) where the dependent variable is the change taken from 1972-1979 in log imports in columns 1-4, change in import penetration, which is imports divided by shipments minus exports plus imports, in column 5, change in log exports in column 6, and change in exports scaled by shipments in column 7. To facilitate comparison with primary estimates, these changes have been made into 9-year equivalent changes. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented in columns 2-7 with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data from Schott (2008) and the NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta \ln(1 + AVE)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.4: Tariff Liberalization and Trade – Column 2 IV

	(1) $\Delta\ln(Imp_i)$	(2) $\Delta\ln(Imp_i)$	(3) $\Delta\ln(Imp_i)$	(4) $\Delta\ln(Imp_i)$	(5) $\Delta ImpPen_i$	(6) $\Delta\ln(Exp_i)$	(7) $\Delta\frac{Exp_i}{Ship_i}$
$\Delta\ln(1+AVE_i)$	-1.683	-8.714***	-9.386***	-8.486***	-1.567***	-3.586	-0.355
	(3.480)	(2.527)	(2.859)	(3.197)	(0.332)	(2.182)	(0.275)
$\Delta\ln(p^*_{i,t-1})$			-0.259**	-0.254**	0.003	-0.153	-0.021
			(0.126)	(0.127)	(0.016)	(0.124)	(0.025)
$STS_i * \frac{AVE_i}{1+AVE_i}$			-1.937	-1.192	-0.847***	1.661	-0.082
			(1.776)	(1.961)	(0.196)	(1.440)	(0.095)
$I(MFA_i)$				-0.040	0.017	-0.091	-0.013*
				(0.086)	(0.015)	(0.074)	(0.007)
$\Delta AVE_i{}^{Exports}$				-0.269	0.025	0.076	0.003
				(0.253)	(0.017)	(0.091)	(0.010)
Constant	1.122***	0.969***	0.984***	0.930***	0.042***	0.445***	0.009
	(0.105)	(0.082)	(0.084)	(0.106)	(0.008)	(0.051)	(0.006)
IQE	-0.049	-0.256	-0.275	-0.249	-0.046	-0.105	-0.010
Estimator	OLS	IV	IV	IV	IV	IV	IV
Obs.	392	392	392	392	392	392	392
1^{st} Stage Coeff.		-0.093	-0.094	-0.089	-0.089	-0.089	-0.089
KP F-Stat	.	169.781	143.742	110.818	110.818	110.818	110.818

Notes: Table presents estimates of equation (1) where the dependent variable is the change taken from 1979-1988 in log imports in columns 1-4, change in import penetration, which is imports divided by shipments minus exports plus imports, in column 5, change in log exports in column 6, and change in exports scaled by shipments in column 7. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented in columns 2-7 with the 1978 column 2 AVE tariffs. Outcome data from Schott (2008) and the NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta\ln(1+AVE)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.5: Tariff Liberalization and Trade – 1930 IV

	(1) $\Delta\ln(Imp_i)$	(2) $\Delta\ln(Imp_i)$	(3) $\Delta\ln(Imp_i)$	(4) $\Delta\ln(Imp_i)$	(5) $\Delta ImpPen_i$	(6) $\Delta\ln(Exp_i)$	(7) $\Delta\frac{Exp_i}{Ship_i}$
$\Delta\ln(1+AVE_i)$	-0.942	-11.138***	-13.307***	-13.088*	-1.991***	-4.598	0.070
	(4.466)	(3.889)	(4.131)	(7.099)	(0.759)	(4.117)	(0.446)
$\Delta\ln(p^*_{i,t-1})$			-0.171	-0.142	0.010	-0.155	-0.031
			(0.141)	(0.183)	(0.021)	(0.138)	(0.028)
$STS_i * \frac{AVE_i}{1+AVE_i}$			-4.139**	-3.455	-1.024***	1.344	0.041
			(1.893)	(2.527)	(0.316)	(1.771)	(0.145)
$I(MFA_i)$				-0.086	0.010	-0.101	-0.010
				(0.163)	(0.020)	(0.098)	(0.010)
$\Delta AVE_i^{Exports}$				-0.237	0.025	0.111	-0.005
				(0.334)	(0.022)	(0.101)	(0.013)
Constant	1.117***	0.909***	0.912***	0.865***	0.042***	0.443***	0.017**
	(0.127)	(0.081)	(0.083)	(0.100)	(0.010)	(0.070)	(0.008)
IQE	-0.038	-0.313	-0.374	-0.368	-0.056	-0.129	0.002
Estimator	OLS	IV	IV	IV	IV	IV	IV
Obs.	311	311	311	311	311	311	311
1^{st} Stage Coeff.		0.104	0.094	0.072	0.072	0.072	0.072
KP F-Stat	.	114.998	76.480	46.591	46.591	46.591	46.591

Notes: Table presents estimates of equation (1) where the dependent variable is a change in log imports in columns 1-4, change in import penetration, which is imports divided by shipments minus exports plus imports, in column 5, change in log exports in column 6, and change in exports scaled by shipments in column 7. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented in columns 2-7 with the Swiss formula implied tariff change as in equations (3) and (4) but replacing the 1978 column 2 tariffs with the 1930 Smoot-Hawley tariffs. Outcome data from Schott (2008) and the NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta\ln(1+AVE)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.6: Tariff Liberalization and Pay Inequality – Column 2 IV

	$\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}}\right)$			
	(1)	(2)	(3)	(4)
$\Delta \ln (1 + AVE_i)$	-1.772**	-2.945***	-2.763***	-2.869***
	(0.693)	(0.826)	(1.011)	(1.018)
IQE	-0.052	-0.087	-0.081	-0.085
Estimator	IV	IV	IV	IV
Controls	None	+Trade	+Prod.	+Tech.
Obs.	395	395	395	395
1^{st} Stage Coeff.	-0.093	-0.090	-0.085	-0.084
KP F-Stat	171.913	113.820	88.746	86.666

Notes: Table presents estimates of equation (1) where the dependent variable is the change in log pay inequality. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented in columns 2-7 with the 1978 column 2 AVE tariffs. Column 1 has no controls. Column 2 adds the trade controls from column 2 of Table 3. Column 3 adds the production controls from column 3 of Table 3. Column 4 adds the technology controls from column 4 of Table 3. Outcome data from Schott (2008) and the NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta \ln(1 + AVE)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.7: Tariff Liberalization and Trade – Controlling for Lagged Changes

	(1) $\Delta\ln(Imp_i)$	(2) $\Delta\ln(Imp_i)$	(3) $\Delta\ln(Imp_i)$	(4) $\Delta\ln(Imp_i)$	(5) $\Delta ImpPen_i$	(6) $\Delta\ln(Exp_i)$	(7) $\Delta\frac{Exp_i}{Ship_i}$
$\Delta\ln(1+AVE_i)$	-1.683	-9.529***	-10.796***	-10.111***	-1.527***	-3.340	-0.294
	(3.480)	(2.202)	(2.361)	(2.840)	(0.368)	(2.190)	(0.258)
$\Delta\ln(Imports_{i,t-1})$		-0.118*	-0.148**	-0.144**			
		(0.069)	(0.072)	(0.071)			
ΔImport Penetration$_{i,t-1}$					0.123		
					(0.237)		
$\Delta\ln(Exports_{i,t-1})$						-0.078*	
						(0.044)	
$\Delta\frac{Exports_i}{Shipments_i}{}_{t-1}$							-0.195***
							(0.071)
IQE	-0.071	-0.280	-0.317	-0.297	-0.045	-0.098	-0.009
Estimator	OLS	IV	IV	IV	IV	IV	IV
Obs.	392	389	389	389	392	390	392
1^{st} Stage Coeff.		0.148	0.148	0.141	0.143	0.142	0.143
KP F-Stat	.	175.892	150.757	113.460	116.828	110.841	115.170

Notes: Table presents estimates of equation (1) where the dependent variable is a change taken from 1979-1988 in log imports in columns 1-4, change in import penetration, which is imports divided by shipments minus exports plus imports, in column 5, change in log exports in column 6, and change in exports scaled by shipments in column 7. All specifications use all trade controls from column 2 of Table 3. As an additional control we include the lagged (1972-1979) change in the dependent variable. To facilitate comparison with primary estimates these lagged changes have been made into 9-year equivalent changes. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented in columns 2-7 with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data from Schott (2008) and the NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta\ln(1+AVE)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.8: Tariff Liberalization and Trade – Upstream Tariffs

	(1) $\Delta\ln(Imp_i)$	(2) $\Delta\ln(Imp_i)$	(3) $\Delta\ln(Imp_i)$	(4) $\Delta\ln(Imp_i)$	(5) $\Delta ImpPen_i$	(6) $\Delta\ln(Exp_i)$	(7) $\Delta\frac{Exp_i}{Ship_i}$
$\Delta\ln(1+AVE_i)$	0.200	-7.458***	-7.905***	-7.000**	-1.069***	-3.183	-0.298
	(3.701)	(2.674)	(3.000)	(3.294)	(0.345)	(2.374)	(0.295)
$\Delta\ln\left(1+AVE_i^{Up}\right)$	-18.461***	-9.681	-11.076	-15.864*	-4.379***	-3.233	-0.310
	(6.503)	(7.312)	(7.326)	(8.269)	(1.249)	(6.743)	(0.596)
$\Delta\ln(p^*_{i,t-1})$			-0.271**	-0.243*	0.005	-0.153	-0.021
			(0.126)	(0.124)	(0.015)	(0.124)	(0.025)
$STS_i * \frac{AVE_i}{1+AVE_i}$			-1.995	-1.102	-0.799***	1.706	-0.073
			(1.771)	(1.986)	(0.195)	(1.462)	(0.097)
$I(MFA_i)$				-0.119	-0.003	-0.106	-0.014*
				(0.095)	(0.015)	(0.074)	(0.007)
$\Delta AVE_i{}^{Exports}$				-0.279	0.021	0.072	0.003
				(0.255)	(0.014)	(0.091)	(0.010)
Constant	1.197***	1.014***	1.038***	1.004***	0.064***	0.462***	0.011
	(0.113)	(0.085)	(0.087)	(0.104)	(0.010)	(0.064)	(0.007)
IQE	0.006	-0.219	-0.232	-0.205	-0.031	-0.093	-0.009
IQE^{up}	-0.542	-0.055	-0.063	-0.090	-0.025	-0.018	-0.002
Estimator	OLS	IV	IV	IV	IV	IV	IV
Obs.	392	392	392	392	392	392	392
1^{st} Stage Coeff.		0.149	0.150	0.143	0.143	0.143	0.143
KP F-Stat	.	74.737	72.269	53.941	53.941	53.941	53.941

Notes: Table presents estimates of equation (1) where the dependent variable is the change in log imports in columns 1-4, change in import penetration, which is imports divided by shipments minus exports plus imports, in column 5, change in log exports in column 6, and change in exports scaled by shipments in column 7. This table also includes upstream tariff changes similarly instrumented. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented in columns 2-7 with the Swiss formula implied tariff change as in equations (3)-(4) and Appendix B.2. Outcome data from Schott (2008) and the NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE and IQE^{up} indicates the effect on the dependent variable of an interquartile increase in $\Delta\ln(1+AVE)$ and $\Delta\ln(1+AVE^{up})$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.9: Tariff Liberalization and Pay Inequality – Decomposition

	(1)	(2)	(3)	(4)
	Panel A: $\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}} \right)_t$			
$\Delta \ln (1 + AVE_i)$	-1.804***	-2.994***	-2.839***	-2.934***
	(0.691)	(0.824)	(1.012)	(1.014)
IQE	-0.053	-0.088	-0.084	-0.086
	Panel B: $\Delta \ln \left(\frac{Emp_i^{Non-Prod}}{Emp_i^{Prod}} \right)_t$			
$\Delta \ln (1 + AVE_i)$	-1.495**	-2.432***	-1.569	-1.681*
	(0.690)	(0.861)	(1.021)	(1.017)
IQE	-0.044	-0.072	-0.046	-0.050
	Panel C: $\Delta \ln \left(\frac{Wage_i^{Non-Prod}}{Wage_i^{Prod}} \right)_t$			
$\Delta \ln (1 + AVE_i)$	-0.309	-0.562	-1.270*	-1.253*
	(0.491)	(0.624)	(0.737)	(0.746)
IQE	-0.009	-0.017	-0.037	-0.037
Estimator	IV	IV	IV	IV
Controls	None	+Trade	+Prod.	+Tech.
Obs.	395	395	395	395
1^{st} Stage Coeff.	0.150	0.144	0.135	0.135
KP F-Stat	180.332	119.386	93.062	91.021

Notes: Table presents estimates of equation (1) where the dependent variable is the change in log pay, log employment, and log wages for non-production workers relative to production workers from 1979-1988 in panels A, B, and C respectively. Column 1 has no controls. Column 2 adds the trade controls from column 2 of Table 3. Column 3 adds the production controls from column 3 of Table 3. Column 4 adds the technology controls from column 4 of Table 3. AVE tariffs have been constructed at the SIC87 level with data from Greenland et al. (2025) and are instrumented with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data from NBER-CES Manufacturing Industry Database (Becker et al., 2021). All data measured at or concorded to the 4-digit 1987 SIC system. First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta \ln(1 + AVE)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.10: Tariff Liberalization and Pay Inequality – Heterogeneity by Gender

	(1)	(2)	(3)	(4)
	Panel A: All Workers			
$\Delta \ln (1 + AVE_i)$	-2.729**	-4.163***	-2.940*	-3.373**
	(1.180)	(1.526)	(1.690)	(1.705)
IQE	-0.075	-0.115	-0.081	-0.093
	Panel B: Male Workers			
$\Delta \ln (1 + AVE_i)$	-1.491	-2.509	-2.072	-2.661
	(1.181)	(1.678)	(1.997)	(1.967)
IQE	-0.041	-0.069	-0.057	-0.074
	Panel C: Female Workers			
$\Delta \ln (1 + AVE_i)$	-9.410***	-10.55***	-7.175***	-6.457***
	(1.937)	(2.228)	(2.419)	(2.174)
IQE	-0.260	-0.292	-0.198	-0.178
Estimator	IV	IV	IV	IV
Controls	None	+Trade	+Prod.	+Tech.
Obs.	72	72	72	72
1^{st} Stage Coeff.	0.138	0.136	0.124	0.131
KP F-Stat	63.995	40.873	32.265	33.857

Notes: Table presents estimates of equation (1). Panel A uses the full sample of workers. Panels B and C restrict the sample to male and female workers respectively. Column 1 has no controls. Column 2 adds the trade controls from column 2 of Table 3. Column 3 adds the production controls from column 3 of Table 3. Column 4 adds the technology controls from column 4 of Table 3. Data measured at or concorded to the Census industry classification. AVE tariffs from Greenland et al. (2025) and are instrumented with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data from IPUMS (Ruggles et al., 2024). First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta \ln(1+AVE_i)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.11: Tariff Liberalization and Pay Inequality – Pretrends by Gender

	(1)	(2)	(3)
	Panel A: $\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}}\right)_{t-1}$		
	All	Men	Women
$\Delta \ln(1 + AVE_i)$	1.591	0.788	2.279
	(3.218)	(3.410)	(3.832)
IQR	0.044	0.022	0.063
Estimator	IV	IV	IV
Obs.	72	72	72
1^{st} Stage Coeff.	0.131	0.131	0.131
KP F-Stat	33.857	33.857	33.857
	Panel B: $\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}}\right)_{t}$		
	All	Men	Women
$\Delta \ln(1 + AVE_i)$	-3.114*	-2.551	-6.632***
	(1.647)	(1.935)	(2.149)
$\Delta \ln \left(\frac{Pay_i^{Non-Prod,A}}{Pay_i^{Prod,A}}\right)_{t-1}$	-0.163**		
	(0.0704)		
$\Delta \ln \left(\frac{Pay_i^{Non-Prod,M}}{Pay_i^{Prod,M}}\right)_{t-1}$		-0.140*	
		(0.0728)	
$\Delta \ln \left(\frac{Pay_i^{Non-Prod,W}}{Pay_i^{Prod,W}}\right)_{t-1}$			0.0769
			(0.0786)
IQR	-0.086	-0.070	-0.183
Estimator	IV	IV	IV
Obs.	72	72	72
1^{st} Stage Coeff.	0.132	0.131	0.134
KP F-Stat	33.236	33.297	34.282

Notes: Table presents estimates of equation (1). Dependent variable is log pay inequality from 1970-1980 in panel A and from 1980-1990 in panel B. Column 1 uses the full sample of workers. Columns 2 and 3 restrict the sample to male and female workers respectively. All specifications use all controls from Table 3. Panel B also includes the change in the outcome variable from 1970-1980 as a control. The superscripts A, M, W indicate the variable relates to all workers, male workers, and women workers. Data measured at or concorded to the Census industry classification. AVE tariffs from Greenland et al. (2025) and are instrumented with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data taken from IPUMS (Ruggles et al., 2024). First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta \ln(1 + AVE_i)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.12: Tariff Liberalization and College Premium – Heterogeneity by Gender

	$\Delta \ln \left(\frac{Pay_i^{Non-Prod}}{Pay_i^{Prod}} \right)_t$		
	(1) All	(2) Men	(3) Women
$\Delta \ln (1 + AVE_i)$	-2.103	-1.449	-6.217
	(2.100)	(2.227)	(3.791)
IQE	-0.058	-0.040	-0.172
Specification	IV	IV	IV
Obs.	72	72	72
1^{st} Stage Coeff.	0.131	0.131	0.131
KP F-Stat	33.857	33.857	33.857

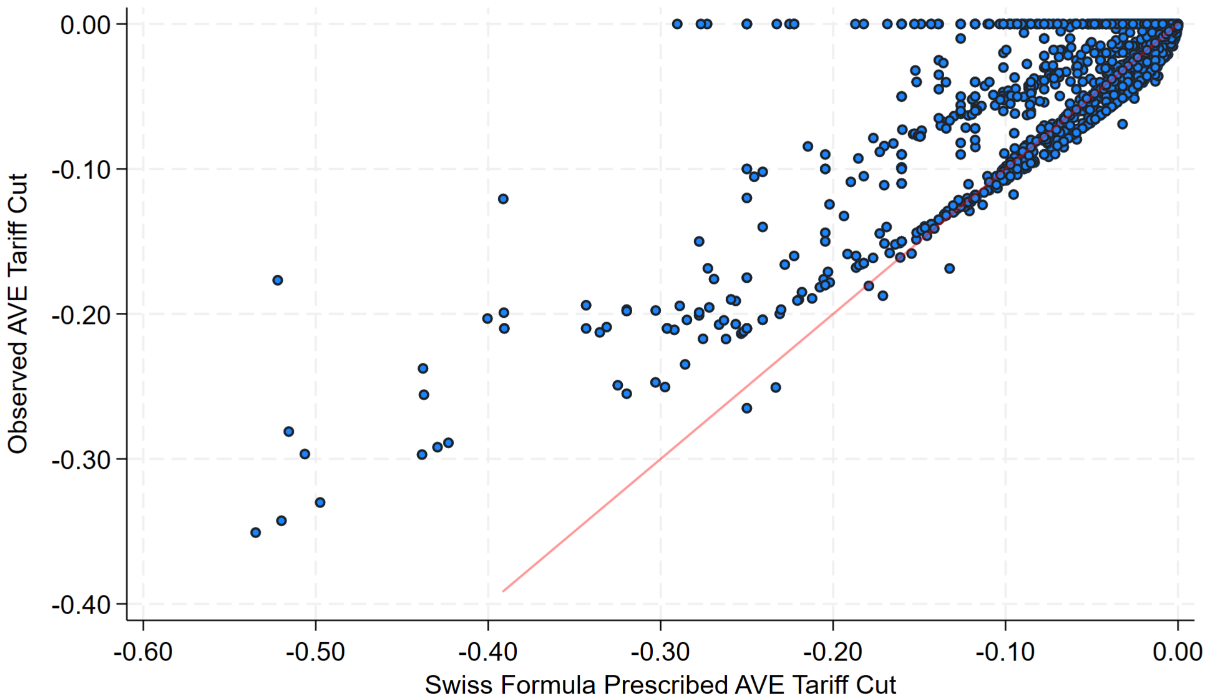
Notes: Table presents estimates of equation (1) where the dependent variable is change in log pay inequality from 1980-1990 among those with a college education relative to those without. Column 1 uses the full sample of workers. Columns 2 and 3 restrict the sample to male and female workers respectively. All columns use all controls from Table 3. Data measured at or concorded to the Census industry classification. AVE tariffs from Greenland et al. (2025) and are instrumented with the Swiss formula implied tariff change as in equations (3) and (4). Outcome data taken from IPUMS (Ruggles et al., 2024). First stage regression coefficients and Kleibergen-Paap Rk F-statistics reported in table footer. IQE indicates the effect on the dependent variable of an interquartile increase in $\Delta \ln(1 + AVE_i)$. Robust standard errors reported in parentheses. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

Table A.13: Occupation Structure by Gender

Occupation	Routineness	Emp Share
	Panel A: Women	
Sewing Machine Operators	0.09	0.11
Electrical Equipment Assemblers	0.92	0.10
Machine Operators	-0.12	0.09
Secretaries	2.54	0.08
Graders and Sorters	0.41	0.06
Bookkeepers	2.02	0.03
Packers and Packagers	0.50	0.03
Production Supervisors	-0.90	0.03
Office Clerks	1.54	0.02
Managers	-1.02	0.02
	Panel B: Men	
Production Supervisors	-0.90	0.09
Machine Operators	-0.12	0.08
Managers	-1.02	0.07
Truck and Tractor Drivers	-1.03	0.04
Electrical Equipment Assemblers	0.92	0.04
Freight and Materials Handlers	-0.24	0.03
Welders and Metal Cutters	0.26	0.03
Machinists	0.37	0.03
Graders and Sorters	0.41	0.03
Industrial Machinery Repairers	-0.34	0.02

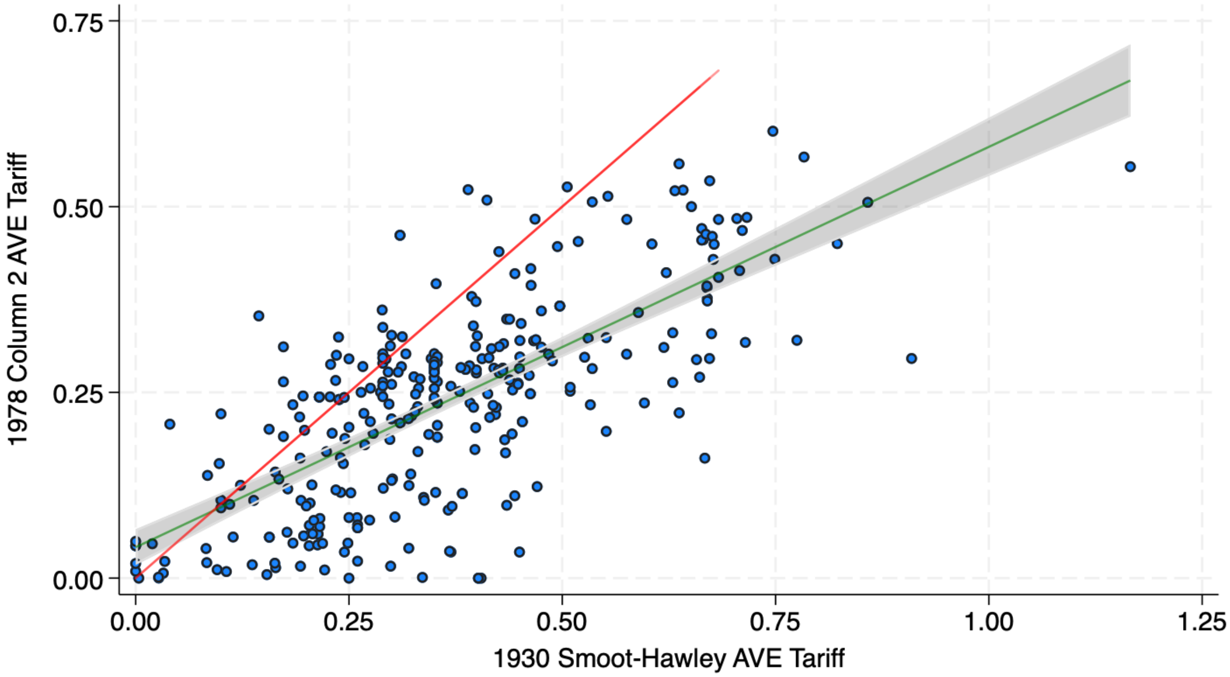
Notes: Table presents the top 10 occupations separately for men (panel A) and women (panel B) according to the routineness index (Autor and Dorn, 2013) and manufacturing employment share. For ease of interpretation in this table, routineness normalized to zero mean and standard deviation equal to 1. Employment and occupation data from IPUMS (Ruggles et al., 2024).

Figure A.1: AVE Tariff Cuts During the Tokyo Round



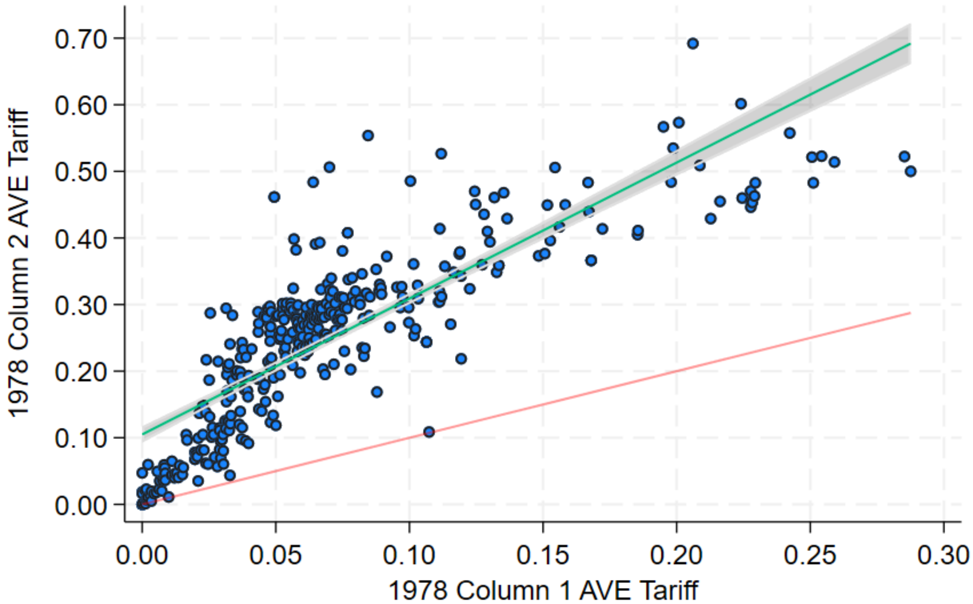
Notes: Figure displays observed AVE tariff cut against the tariff cut prescribed by the Swiss formula at the 5-digit TSUSA level between 1979 and 1988 as described in equation (2). Red line is the 45 degree line which represents strict adherence to the Swiss formula. Tariff data from Greenland et al. (2025).

Figure A.2: 1978 Column 2 vs 1930 AVE Tariffs

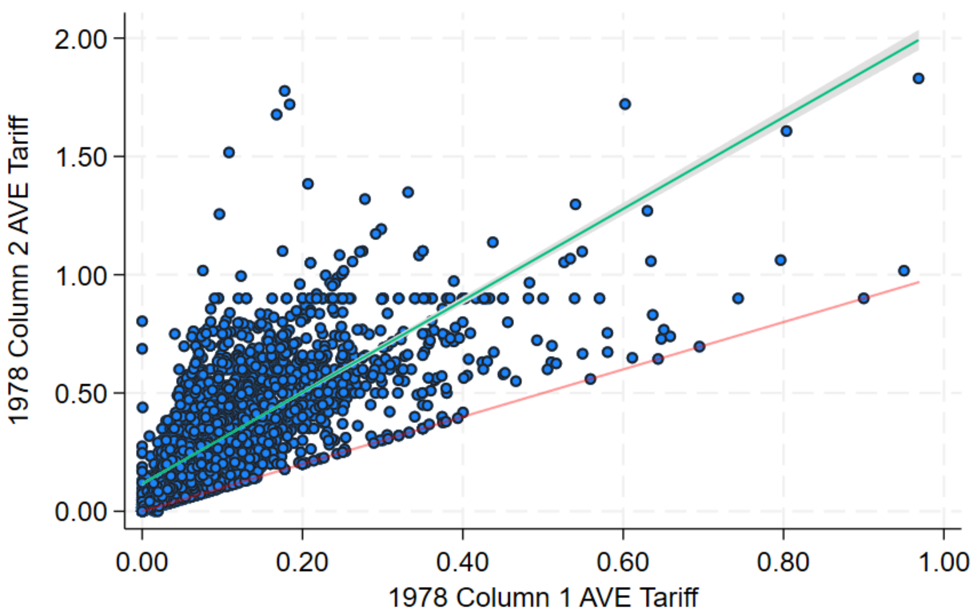


Notes: The figure depicts 1978 column 2 AVE tariffs versus the 1930 AVE tariffs at the 4-digit SIC level. Red line is the 45 degree line. 1978 tariff data from Greenland et al. (2025). 1930 Smoot-Hawley tariff data from Greenland and Lopresti (2024).

Figure A.3: 1978 Column 1 vs 1978 Column 2 AVE Tariffs



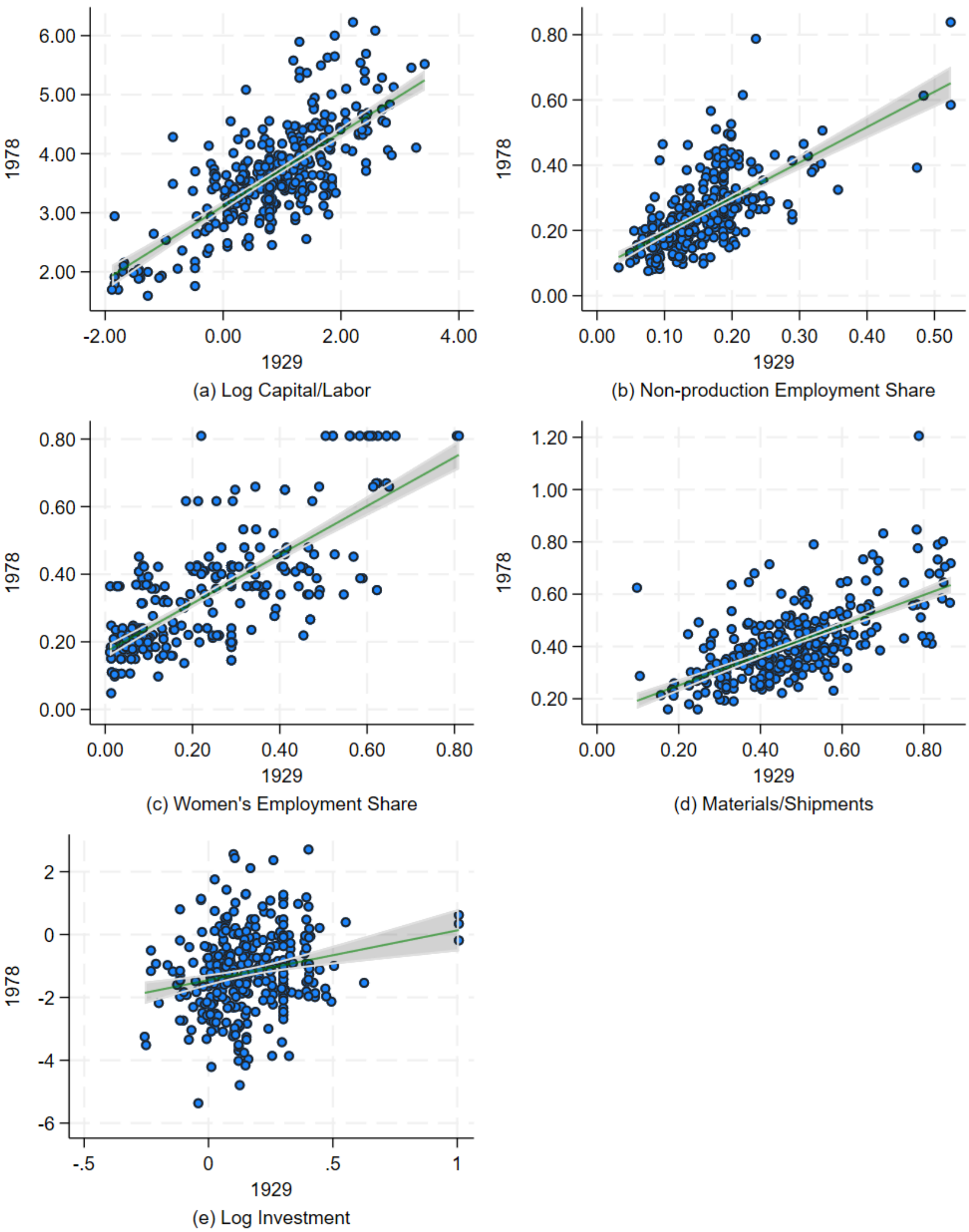
(a) 4-digit SIC Industry Level



(b) 5-digit TSUSA Level

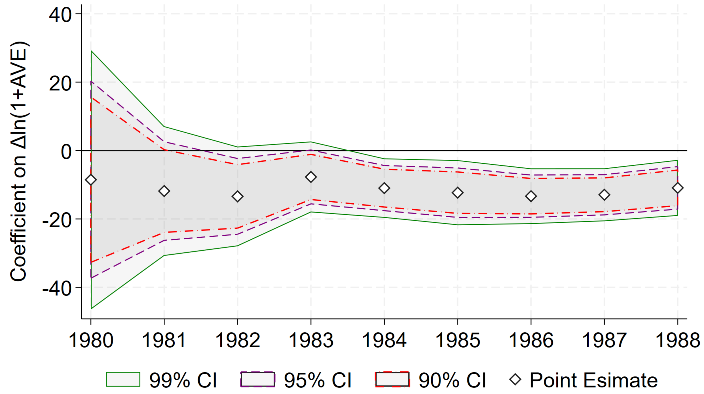
Notes: The figure depicts 1978 Column 1 AVE tariffs versus the 1978 Column 2 AVE tariffs at the 4-digit SIC level in panel (a) and the 5-digit TSUSA level in panel (b). Red line is the 45 degree line. Tariff data from Greenland et al. (2025).

Figure A.4: Persistence of Industry Characteristics 1929-1978

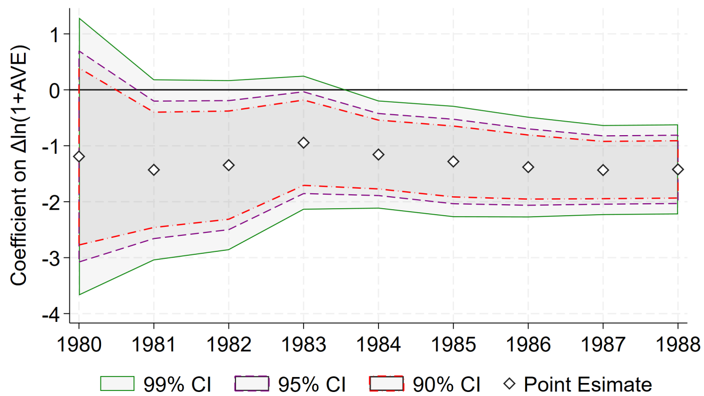


Notes: Panels scatter 1929 and 1978 values of controls. 1929 log capital to labor ratio and log investment defined as, respectively, log of horsepower per worker and the 1927-1929 log change in horsepower. 1978 (1929) Non-production employment share defined as non-production employment (salaried employment) as share of production and non-production (salaried and wage earner) employment. Data from NBER-CES Manufacturing Industry Database (Becker et al., 2021) and 1930 Census of Manufactures.

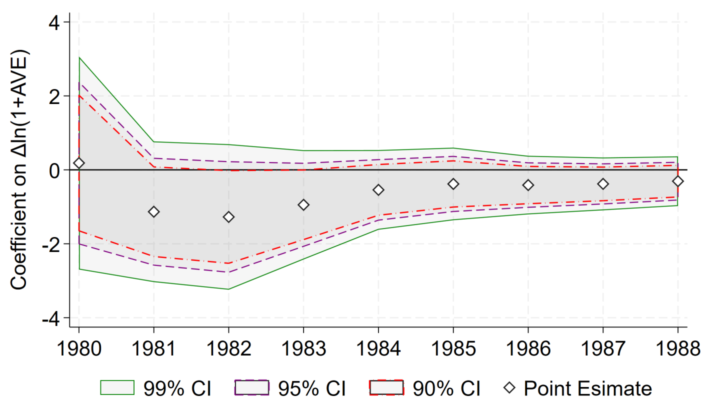
Figure A.5: Tariff Liberalization and Trade – Alternative Time Horizons



(a) $\Delta \ln(Imports)$



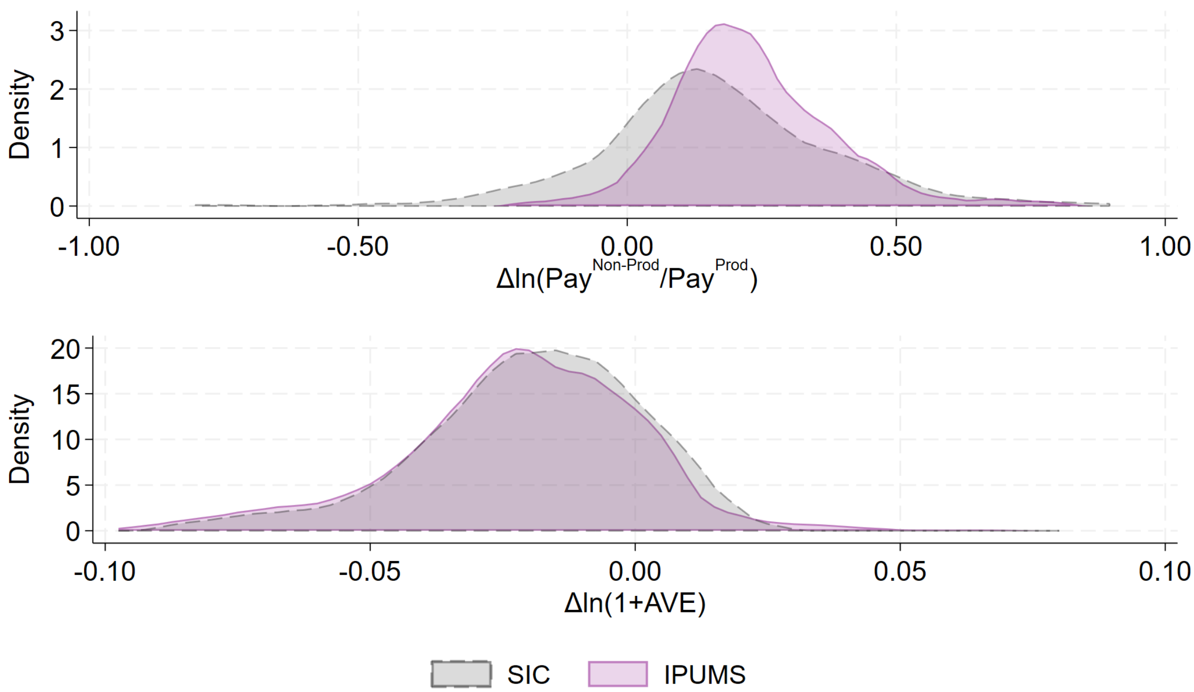
(b) ΔImport Penetration



(c) $\Delta \frac{Exports}{Shipments}$

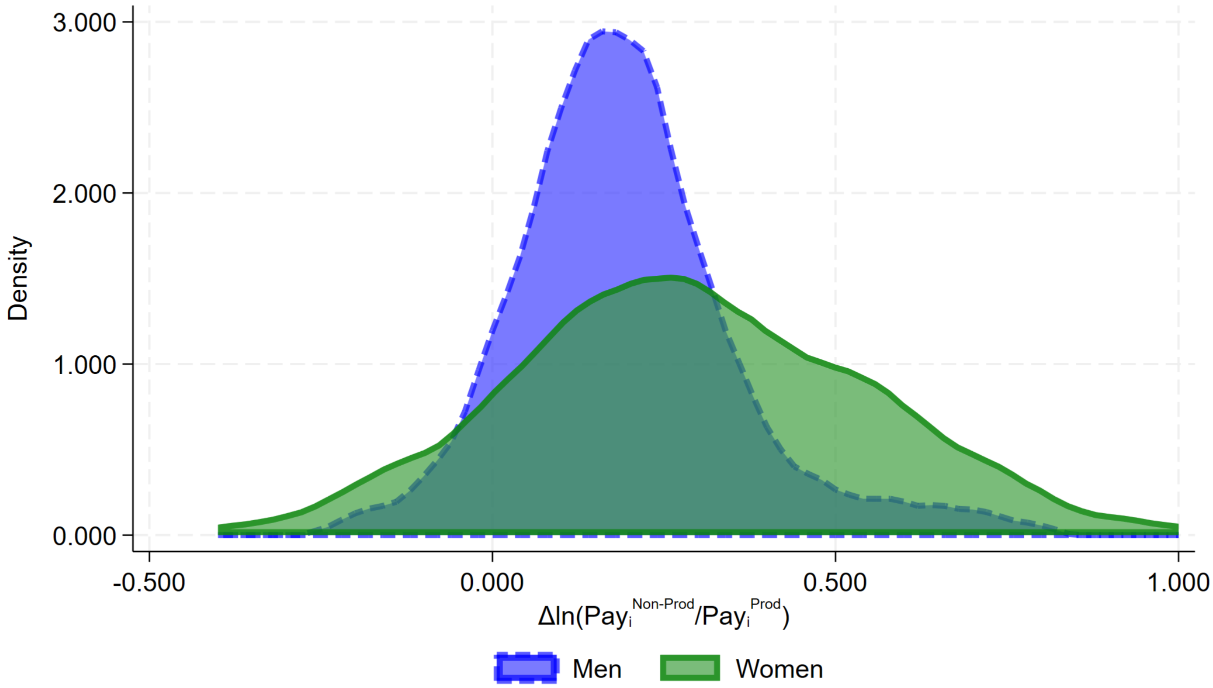
Notes: Figure presents IV estimates from separately estimating equation (1) for each time window between 1979 and the year indicated on the x-axis. Dependent variable listed in panel title. Specifications use full controls as in column 4 of Table 3. Data from Schott (2008), Becker et al. (2021), and Greenland et al. (2025). Confidence intervals reflect robust standard errors.

Figure A.6: Tariff Liberalization and Pay Inequality – Alternative Datasets



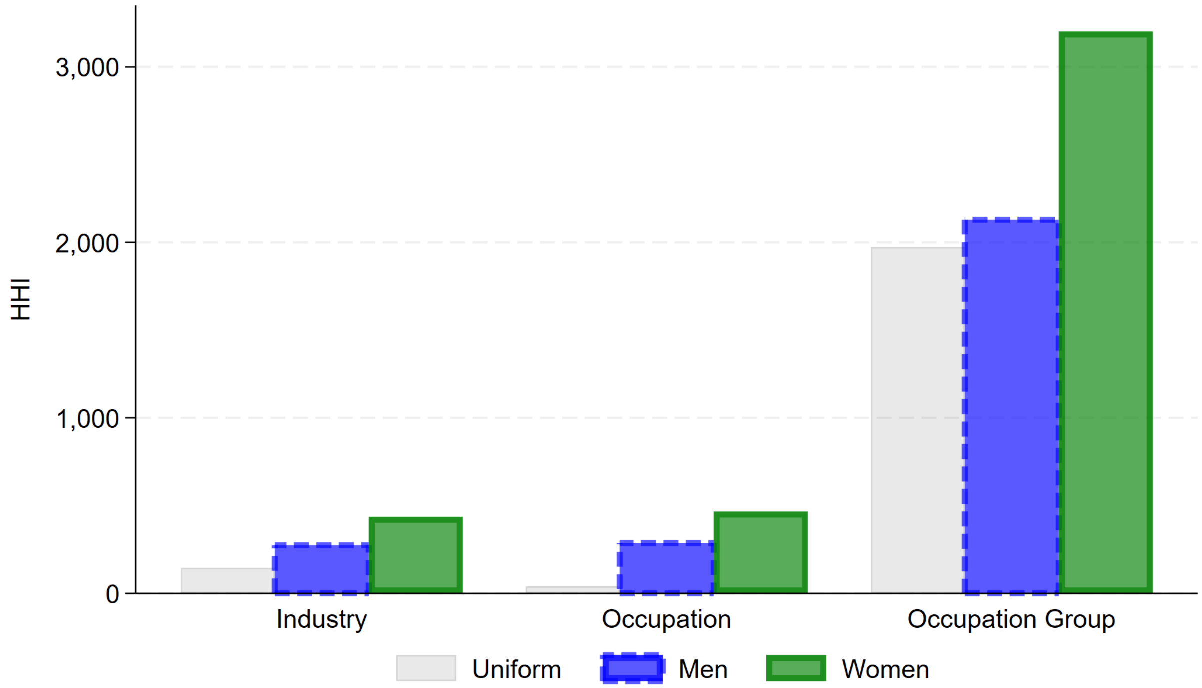
Notes: Figure presents distribution of growth in log pay inequality and change in AVE tariffs ($\Delta\,(\ln(1 + AVE))$ between 1979 and 1988 for the 4-digit SIC87 industries (SIC) and Census industries (IPUMS). Data from Becker et al. (2021), Ruggles et al. (2024), and Greenland et al. (2025).

Figure A.7: Growth in Pay Inequality by Gender



Notes: Figure presents distribution of growth in log pay inequality for men and women between 1980 and 1990 when aggregated to the 72 manufacturing Census industries in our analysis. Data from IPUMS (Ruggles et al., 2024).

Figure A.8: Employment Concentration by Gender



Notes: Figure presents the HHI for men and women by employment at the level of the Census industries, Census occupations, and the occupation groups of Autor and Dorn (2013). Maximum possible HHI is 10,000.

B Data Appendix

B.1 Aggregating Legislated Tariffs to SIC Industry Tariffs

Our goal is to calculate annual AVE tariffs at the 4-digit SIC level, defined as duties collected divided by the value of imports. To do so, we concord from the disaggregate TSUSA level to the 4-digit SITC revision 2 classification and then to the 4-digit SIC 1987 system via the 6-digit HS level. Two issues emerge immediately. First, the concordance from the TSUSA system to the 4-digit SITC revision 2 level maps uniquely from 7-digit TSUSA codes but legislated tariffs are defined at the 5-digit TSUSA level. Second, exporters of a given 7-digit TSUSA good may face one of three tariffs: (i) the vast majority of exporters and products face the column 1 tariffs but, in practice during our sample period, (ii) a small subset of communist countries face column 2 tariffs, and (iii) a subset of products from a subset of developing countries have duty free access (e.g. due to the US Generalized System of Preferences (GSP) program). To this end, we use annual legislated tariffs and import data between 1978 and 1988 from Greenland et al. (2025) at the exporter by 7-digit level.

Starting at the 7-digit TSUSA level, duties collected on good k in year t across exporters n are

$$d_{kt} = \sum_{n} \left(\tau_{knt} m_{knt} + f_{knt} q_{knt}\right) \tag{B.1}$$

where τ is the ad valorem tariff, f is the specific tariff, q is the quantity of imports, and m is the value of imports. Aggregating across the set of 7-digit TSUSA goods $K(j)$ that uniquely map to 4-digit SITC j gives

$$d_{jt} = \sum_{k \in K(j)} d_{kt} \tag{B.2}$$

$$m_{jt} = \sum_{k \in K(j)} m_{kt}. \tag{B.3}$$

For later purposes, it is useful to note that we can always rewrite an aggregate AVE tariff – defined as aggregate duties collected divided by the aggregate value of imports (for example, aggregated across 7-digit TSUSA goods k to the 4-digit SITC industry j) – as an import weighted average of disaggregate AVE tariffs. To this end, let $Col1$ and $Col2$ denote that imports are subject to column 1 and column 2 tariffs. Then,

$$\begin{aligned}
AVE_{jt} \equiv \frac{d_{jt}}{m_{jt}} &= \left[\sum_{k\in K(j)} d_{kt}^{Col1} + \sum_{k\in K(j)} d_{kt}^{Col2}\right]\left[\sum_{k\in K(j)} m_{kt}\right]^{-1} \\
&= \frac{1}{m_{jt}} \sum_{k\in K(j)} \left(\frac{d_{kt}^{Col1}}{m_{kt}^{Col1}} m_{kt}^{Col1} + \frac{d_{kt}^{Col2}}{m_{kt}^{Col2}} m_{kt}^{Col2}\right) \\
&\equiv \frac{1}{m_{jt}} \sum_{k\in K(j)} \left(AVE_{kt}^{Col1} m_{kt}^{Col1} + AVE_{kt}^{Col2} m_{kt}^{Col2}\right) \\
&= \sum_{k\in K(j)} \left(AVE_{kt}^{Col1} \frac{m_{kt}^{Col1}}{m_{kt}} \frac{m_{kt}}{m_{jt}} + AVE_{kt}^{Col2} \frac{m_{kt}^{Col2}}{m_{kt}} \frac{m_{kt}}{m_{jt}}\right) \\
&\equiv \sum_{k\in K(j)} \left(AVE_{kt}^{Col1} \omega_{kt}^{Col1} + AVE_{kt}^{Col2} \omega_{kt}^{Col2}\right) \omega_{kjt} \qquad \text{(B.4)} \\
&\equiv \sum_{k\in K(j)} AVE_{kt} \omega_{kjt}. \qquad \text{(B.5)}
\end{aligned}$$

Equation (B.5) says the aggregate AVE tariff AVE_{jt} is simply the import weighted average of the disaggregate AVE tariffs AVE_{kt} with the weights ω_{kjt} denoting the time-varying share of industry j imports accounted for by good k. And, equation (B.4) says each disaggregate AVE tariff AVE_{kt} is the import weighted average AVE tariff across column 1 imports, AVE_{kt}^{Col1}, and column 2 imports, AVE_{kt}^{Col2}, with the weights ω_{kt}^{Col1} and ω_{kt}^{Col2} denoting, respectively, the time-varying share of good k imports accounted for by column 1 and column 2 imports.[44]

To concord from 4-digit SITC industries j to 4-digit SIC industries i, we use the HS system as an intermediate step. In general, 4-digit SITC codes are more aggregate than 6-digit HS codes. Thus, a 4-digit SITC code j typically maps to multiple 6-digit HS codes. Denote the number of such 6-digit HS codes by l_j^{HS}. Similarly, 6-digit HS6 codes are generally more disaggregate than 4-digit SIC codes. Thus, a 6-digit HS code h typically maps to a unique 4-digit SIC code. Nevertheless, denote the number of such 4-digit SIC codes by l_h^{SIC}. Then the number of "connections" from 4-digit SITC code j to 4-digit SIC code i is $l_j^{HS} l_h^{SIC}$. Naturally, some of these connections start with the same 4-digit SITC code j and end with the same 4-digit SIC code i. Thus, denote the number of such unique 4-digit SIC codes "connected" from 4-digit SITC code j by l_j^{SIC}.[45]

[44]Equation (B.4) omits a third term for the category of imports subject to zero tariffs due to, e.g. the US GSP program. This is without loss of generality because these imports have zero AVE tariffs and still enter the total import terms m_{kt} and m_{jt}.

[45]For example, SITC code 0011 maps to the two HS codes 010210 and 010290 (i.e. $l_j^{HS} = 2$). And, each

Let $J(i)$ denote the set of SITC industries j that concord to SIC industry i, and apportion SITC industry j's duties and imports equally across connected SIC industries i. Then, the AVE tariff for 4-digit SIC industry i is the weighted sum of duties collected across 4-digit SITC industries $j \in J(i)$ divided by the weighted sum of imports across these 4-digit SITC industries:

$$\begin{aligned} AVE_{it} &= \left[\sum_{j \in J(i)} \frac{1}{l_j^{SIC}} d_{jt}\right]\left[\sum_{j \in J(i)} \frac{1}{l_j^{SIC}} m_{jt}\right]^{-1} \\ &\equiv \left[\sum_{j \in J(i)} d_{jit}\right]\left[\sum_{j \in J(i)} m_{jit}\right]^{-1} \\ &= \sum_{j \in J(i)} AVE_{jt}\omega_{jit} \end{aligned} \tag{B.6}$$

where the last line follows from the general interpretation of an aggregate AVE in equation (B.5). Equation (B.6) says the AVE tariff for 4-digit SIC industry i, AVE_{it}, is the import-weighted average of AVEs across the associated 4-digit SITC industries j, AVE_{jt}, where the weights $\omega_{jit} = \frac{m_{ji}}{\sum_j m_{ji}}$ correspond to the time varying share of industry i's imports accounted for by industry j. And, as described above, equation (B.5) says the AVE tariff for SITC industry industry j is the import-weighted average AVE tariff across 7-digit TSUSA goods k which, per equation (B.4), is the import weighted average AVE tariff across good k column 1 and column 2 imports. Thus, ultimately, the AVE tariff for a 4-digit SIC industry i is an import weighted average of the AVE of the underlying legislated tariffs at the TSUSA tariff-line level.

Construction of our instruments now follows naturally in two steps. First, we calculate an instrumented value for AVE_{kt} with fixed 1978 import weights to substitute into equation (B.5). This gives instrumented values for AVE_{jt}. Second, we substitute the instrumented values for AVE_{jt} into equation (B.6) using fixed 1978 import weights. This gives instrumented values for AVE_{it}.

Our Swiss IV substitutes column 2 tariffs into the 1978 Swiss formula and phases out linearly between 1980 and 1987. Note that Appendix Figure A.1 re-creates Figure 4 at the tariff line 5-digit TSUSA level. And, panel A of Appendix Figure A.3 plots the column 1

of these HS codes map to the two SIC codes 0211 and 0241 (i.e. $l_h^{SIC} = 2$ for each of the two HS codes). But, among these $l_j^{HS} l_h^{SIC} = 4$ connections, the SITC code 0011 only maps to two SIC codes (i.e. l_j^{SIC}=2): 0211 and 0241.

1978 AVE tariff against the 1978 column 2 AVE tariff at the tariff-line 5-digit TSUSA level. In 1979, the instrument value for good g at the 5-digit TSUSA level is

$$
\begin{aligned}
AVE^{IV}_{g,1979} &= \frac{\sum_n \left(\tau^{Col2}_{gn,1978} m^{Col1}_{gn,1978} + f^{Col2}_{gn,1978} q^{Col1}_{gn,1978}\right)}{m^{Col1}_{g,1978}} \\
&= \tau^{Col2}_{g,1978} + \frac{f^{Col2}_{g,1978}}{p_{g,1978}}
\end{aligned}
\tag{B.7}
$$

where $p_{g,1978}$ is the 1978 unit value for 5-digit TSUSA good g column 1 imports. Once the US has fully implemented their Tokyo tariff cuts in $t = 1987, 1988$, the instrument value is:

$$
AVE^{IV}_{gt} = \frac{0.14 \times AVE^{IV}_{g,1979}}{0.14 + AVE^{IV}_{g,1979}} \text{ for } t = 1987, 1988
\tag{B.8}
$$

For the phase-out years $t = 1980, ..., 1986$, we phase out AVE^{IV}_{gt} linearly between $AVE^{IV}_{g,1979}$ and $AVE^{IV}_{g,1987}$. Substituting AVE^{IV}_{gt} for AVE^{col1}_{gt} in equation (B.4) and fixing import weights at their 1978 values gives:

$$
\begin{aligned}
AVE^{IV}_{jt} &\equiv \sum_{k \in K(j)} \left(AVE^{IV}_{g(k),t} \omega^{Col1}_{k,1978} + AVE^{Col2}_{g(k),1978} \omega^{Col2}_{k,1978}\right) \omega_{kj,1978} \\
&= \sum_{g \in G(j)} \left(AVE^{IV}_{gt} \omega^{Col1}_{g,1978} + AVE^{Col2}_{g,1978} \omega^{Col2}_{g,1978}\right) \omega_{gj,1978} \\
&\equiv \sum_{g \in G(j)} AVE^{IV}_{gt} \omega_{gj,1978}.
\end{aligned}
\tag{B.9}
$$

where $g(k)$ in the first line denotes the 5-digit TSUSA good g associated with 7-digit TSUSA good k. Finally, substituting equation (B.9) into equation (B.6) gives

$$
\begin{aligned}
AVE^{IV}_{it} &\equiv \sum_{j \in J(i)} AVE^{IV}_{jt} \omega_{ji,1978} \\
&= \sum_{j \in J(i)} \sum_{g \in G(j)} AVE^{IV}_{gt} \omega_{gj,1978} \omega_{ji,1978}.
\end{aligned}
\tag{B.10}
$$

That is, the value of our Swiss IV for 4-digit SIC industry i is an import weighted average of the Swiss IV at the tariff line, i.e. 5-digit TSUSA, level g.

B.2 Upstream Input Tariffs

Our measure of input tariffs faced by industry i is a weighted average of AVE tariffs on all manufacturing inputs z used by industry i. The time-invariant weights depend on (i) s_{iz}, the total requirement for industry i of input z per dollar of industry i output from the 1977 IO table from the US Bureau of Economic Analysis, and (ii) 1978 US imports of z, m_z, as a share of 1978 US manufacturing imports (Schott, 2008). Specifically,

$$AVE_{it}^{Up} = \sum_{z} AVE_{zt}\omega_z \tag{B.11}$$

where the time-invariant weight is

$$\omega_z = \frac{s_{iz}\frac{m_z}{\sum_z m_z}}{\sum_z s_{iz}\frac{m_z}{\sum_z m_z}}. \tag{B.12}$$

Finally, using equation (B.10) to substitute AVE_{zt}^{IV} into (B.11) for AVE_{zt} gives our instrument for AVE_{it}^{Up}. Note that, for industry i, the weight on its input tariff z relative to the weight on its input tariff z' is

$$\frac{\omega_z}{\omega_{z'}} = \frac{s_{iz}}{s_{iz'}}\frac{m_z}{m_{z'}}. \tag{B.13}$$

Intuitively, this relative weight is higher when (i) industry i uses more of input z relative to input z' and/or (ii) the US imports more of input z than input z'.

B.3 Tariff Changes by Foreign Countries

We digitize data on the negotaited Tokyo Round tariff concessions granted by Japan, the European Economic Community (EEC) and Canada using data from the WTO website.[46]. These are the primary export destinations for US manufacturing, covering over half of US exports in 1979. In Appendix Figure B.1 we show that, apart from Mexico, no more than 3% of US exports go to any other country. And, 129 countries constitute the remaining 17% of US manufacturing export value that goes to the rest of the world.

Figure B.1 about here.

We concord these tariff lines to the 4-digit SIC level. For Japan and the EEC, we concord to the Customs Co-operation Council Nomenclature (CCCN) and then to 1992 HS codes

[46] https://www.wto.org/english/docs_e/gattbilaterals_e/indexbyround_e.htm

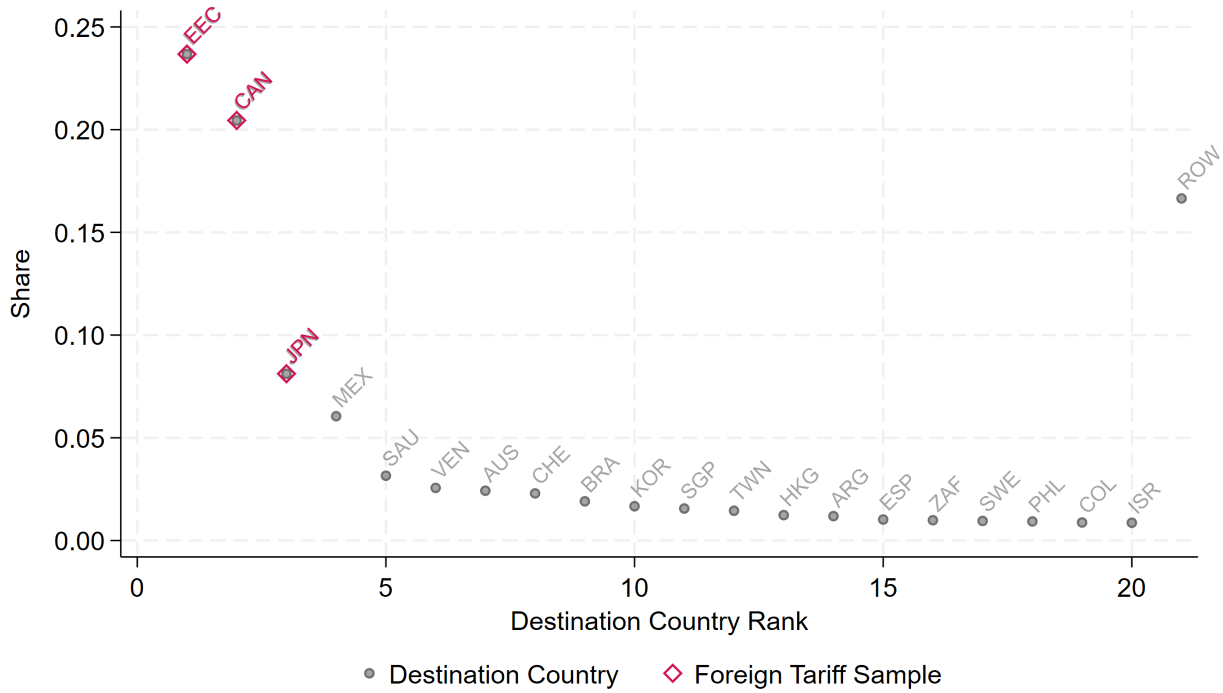
using a concordance provided by WITS.[47]. Finally, we concord these to the SIC using the concordance created by Pierce and Schott (2012). For Canada, we concord from the Canadian SIC system to the US SIC system using data and a concordance from Trefler (2004). Finally, we weight these partner-specific SIC-level tariff changes by the share of 1972 US exports among the three partners accounted for by each partner.

The WTO/GATT tariff data only provide information on tariff cuts. They do not provide information on tariffs that do not change. Thus, we cannot construct $\Delta \ln(1 + AVE_i^{Exports})$ since it requires the initial tariffs. Hence, we use $\Delta AVE_i^{Exports}$ which does not require initial tariffs. This measure is correlated with US tariff cuts, as seen in Appendix Figure B.2.

Figure B.2 about here.

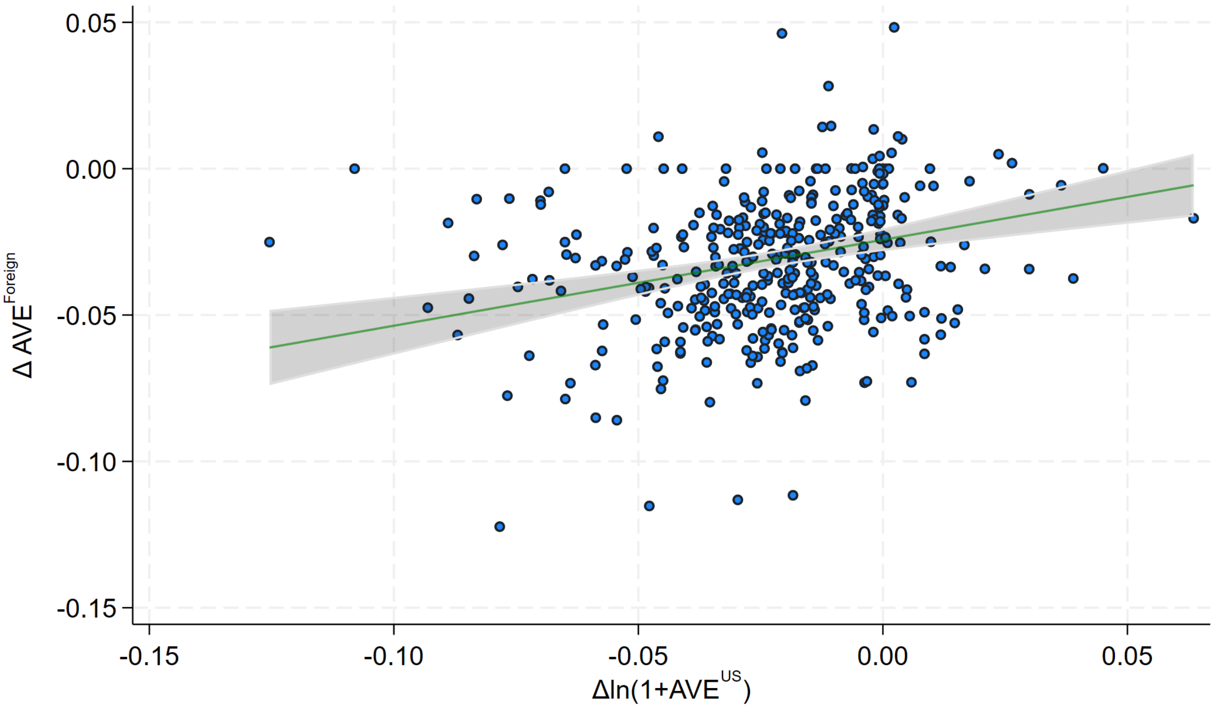
[47] https://wits.worldbank.org/product_concordance.html

Figure B.1: Destination Country Share of 1979 US Exports



Notes: Figure displays the share of 1979 US manufacturing exports to various destination countries. Foreign tariff sample used in main text includes EEC, Canada, and Japan. Data from Schott (2008).

Figure B.2: US and Foreign Country Tariff Cuts During Tokyo Round of the GATT



Notes: Figure scatters US tariff cuts $\Delta\ln\left(1 + AVE^{US}\right)$ against export weighted average foreign country tariff cuts $\Delta AVE^{Foreign}$ by EEC, Japan, and Canada during the Tokyo Round using 1972 US export weights. One observation (driven by a large increase in tariffs related to malt and brewing in Canada) omitted for readability. Tariff data digitized from `https://www.wto.org/english/docs_e/gattbilaterals_e/indexbyround_e.htm` for EEC and Japan and taken taken from Trefler (2004) for Canada.

C Endogenous Liberalization

An appealing feature of our setting is that we can explore the nature of endogenous protection: what factors drove US policymakers to deviate from the Swiss formula? Specifically, we explore the characteristics of the under-liberalized industries relative to the over-liberalized industries. This is a rare opportunity, as studies analyzing the causal impact of trade policy typically require settings in which the observed policy changes are plausibly exogenous – ruling out an exploration of the endogenous determinants of those changes by definition – or require that all such determinants have been identified. By exploiting the planned tariff reductions from the Swiss formula, we can identify the causal effects of Tokyo Round tariff reductions while simultaneously identifying endogenous correlates of observed protection.

Figure C.1 about here.

Motivated by our focus on income inequality between production and non-production workers, we begin by highlighting the imprint of industry skill intensity on the Tokyo Round tariff cuts chosen by policymakers. Appendix Figure C.1 presents separate kernel densities of the 1978 non-production labor share of employment within the 198 under-liberalized (blue) industries and the 197 over-liberalized industries (red). While there is considerable overlap in the densities, the overall pattern is stark: under-liberalized industries tend to have smaller employment shares of non-production labor than over-liberalized industries. That is, policymakers are more likely to deviate from the Swiss formula to protect industries with lower pre-Tokyo skill intensity.

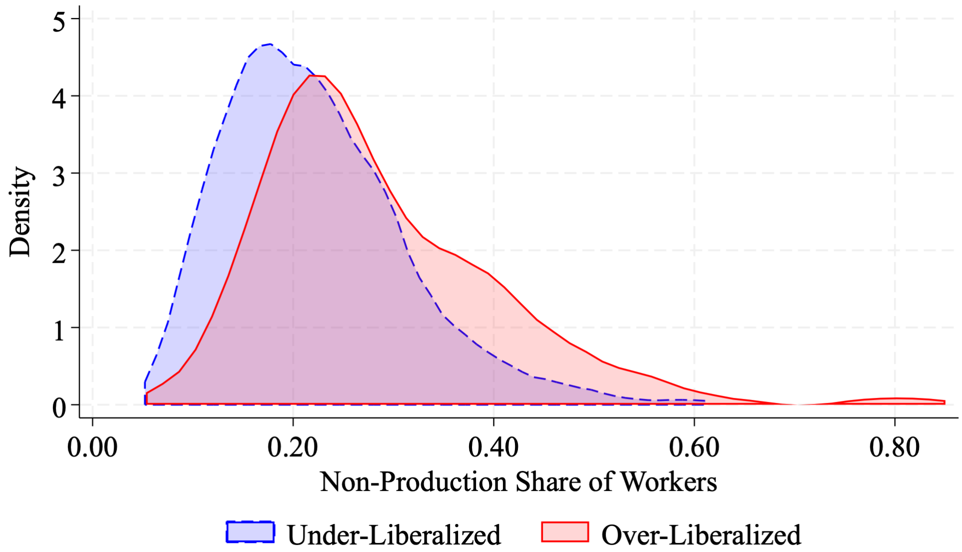
We analyze this relationship more formally in Appendix Table C.1. Specifically, we estimate a linear probability model relating 1978 industry characteristics to whether or not an industry is under-liberalized. Column (1) of the table includes only log of one plus 1978 observed industry tariff levels. Column 2 introduces our trade controls including tariffs by foreign countries, lagged industry price growth, a dummy variable for having MFA exposure, and a control for the inflationary erosion of specific tariffs. Column 3 introduces capital to labor and women's employment shares. Column 4 introduces non-production employment shares. Finally, column 5 introduces our technology-related controls.

Three characteristics predict under-liberalization. First, industries experiencing substantial lagged price growth are more likely to be under-liberalized (have their tariffs cut by less than prescribed by the Swiss formula). As documented in Greenland et al. (2025), the years leading up to the Tokyo round experienced rampant inflation which eroded the protective capacity of specific tariffs. As such, many of the sectors experiencing industry price growth were liberalized during the 1970s. These industries seem to have been partially exempted from the full tariff cuts. Second, industries whose labor forces were tilted toward production workers were also differentially protected. This is consistent with our findings in Appendix Figure C.1. Lastly, industries with large concentrations of occupations heavily characterized by routine tasks are less likely to have been exempted from the full tariff cuts implied by the Swiss formula.

Table C.1 about here.

These relationships suggest that politicians considered pre-existing characteristics in a way that materially impacted the distribution of Tokyo liberalization across industries. Naturally, such endogenous manipulation of tariff changes may yield biased estimates of the effects of observed changes on industry outcomes. This motivates the IV approach taken in the paper.

Figure C.1: Non-production Labor Shares and Endogenous Liberalization



Notes: Kernel density of non-production labor shares of employment displayed separately for 4-digit 1987 SIC industries receiving lower tariff cuts than prescribed by the Swiss formula (Under-Liberalized) versus those receiving greater tariff cuts than prescribed (Over-Liberalized). Data from Greenland et al. (2025), the NBER-CES Manufacturing Industry Database (Becker et al., 2021).

Table C.1: Determinants of Under-Liberalization

	(1)	(2)	(3)	(4)
$\ln(1 + AVE_i)$	2.804***	1.135*	0.833	1.240
	(0.366)	(0.649)	(0.774)	(0.769)
$\Delta AVE_i^{Exports}$		0.007	0.019	-0.052
		(0.095)	(0.100)	(0.092)
$\Delta \ln(p^*_{i,t-1})$		0.445***	0.411***	0.321**
		(0.136)	(0.130)	(0.128)
$I(MFA_i)$		0.223***	0.161**	0.139*
		(0.075)	(0.079)	(0.079)
$STS_i * \frac{AVE_i}{1+AVE_i}$		2.925**	2.040	1.202
		(1.156)	(1.237)	(1.271)
$\frac{Capital_i}{Labor_i}$			0.047	0.036
			(0.039)	(0.045)
$\frac{Emp_i^{Non-Prod}}{Emp_i}$			-0.904***	-0.778***
			(0.206)	(0.221)
$\frac{Emp_i^{Women}}{Emp_i}$			0.193	0.175
			(0.212)	(0.222)
$\ln(Investment_i)$				-0.007
				(0.022)
$\Delta \ln(Investment_{i,t-1})$				0.117***
				(0.044)
$Routineness_i$				-0.212**
				(0.105)
$Automation_i$				-0.092
				(0.057)
$\frac{Materials_i}{Shipments_i}$				0.270
				(0.188)
Constant	0.318***	0.335***	0.373*	0.548*
	(0.039)	(0.045)	(0.202)	(0.281)
Obs.	395	395	395	395
Adj. R^2	0.075	0.130	0.161	0.189

Notes: Dependent variable is a dummy variable taking on the value of 1 if an industry's observed AVE tariff cut was smaller than prescribed by the Swiss formula. OLS estimation of linear probability models. Robust standard errors. *, **, and *** indicate p-values less than 10%, 5% and 1%, respectively.

D Berman et al. (1994) Estimates

Consider groups denoted by $i = 1, ..., I$ and types denoted by $j = 1, ..., J$. In our context, groups are industries and types are production and non-production workers. Let y_{ji} denote the "within" variable of interest for type j within group i. In our context, this is the non-production (NP) pay share within industry i:

$$y_{ji} = \frac{pay_{NP,i}}{pay_i} \tag{B.14}$$

Let y_i denote the "between" shares for group i. In our setting, this is industry i's share of total pay:

$$y_i = \frac{pay_i}{\sum_i pay_i} = \frac{pay_i}{pay}. \tag{B.15}$$

We are interested in how the "aggregate" y_j, in our context the economy-wide non-production pay share y_{NP}, changes over time. More specifically, the variable of interest is

$$\begin{aligned}
\Delta y_j \equiv \Delta y_{NP} &= \Delta \frac{pay_{NP}}{pay} \\
&= \frac{\sum_i pay_{NP,i,t}}{pay_t} - \frac{\sum_i pay_{NP,i,t-1}}{pay_{t-1}} \\
&= \sum_i \frac{pay_{NP,i,t}}{pay_{it}} \frac{pay_{it}}{pay_t} - \sum_i \frac{pay_{NP,i,t-1}}{pay_{i,t-1}} \frac{pay_{i,t-1}}{pay_{t-1}} \\
&= \sum_i y_{NP,it} y_{it} - \sum_i y_{NP,i,t-1} y_{i,t-1} \qquad \text{(B.16)}
\end{aligned}$$

This can be decomposed into "within" and "between" components

$$\begin{aligned}
\Delta y_{NP} &= \sum_i \left(y_{NP,i,t-1} + \frac{1}{2} \Delta y_{NP,i} \right) \Delta y_i + \sum_i \left(y_{i,t-1} + \frac{1}{2} \Delta y_i \right) \Delta y_{NP,i} \qquad \text{(B.17)} \\
&= \sum_i \bar{y}_{NP,i} \Delta y_i + \sum_i \bar{y}_i \Delta y_{NP,i}. \qquad \text{(B.18)}
\end{aligned}$$

The first summation term is the between component, driven by the between-industry change in pay shares weighted by the industry-specific non-production pay share, with weights based on an average over time. The second summation term is the within component, driven by the change in the within-industry non-production pay shares weighted by industry pay shares, with weights based on an average over time.

Our aim is to obtain estimates of Δy_{NP} caused by tariff changes in order to calculate the share of the observed change in y_{NP} due to tariffs through both the within and between components. In equation (B.18), the $t-1$ variables are observable and we can estimate Δy_i and $\Delta y_{NP,i}$.

Our main analysis focuses on estimating the change in within-industry non-production pay shares $\Delta y_{NP,i}$ but does not address estimation of the between-industry pay shares. An important complication in estimating these is that they must sum to 1, something which does not constrain estimation of within-industry non-production pay shares. To address this, we estimate industry-level changes in pay and use the set of predicted changes to calculate the predicted change in industry pay shares. That is

$$\begin{aligned}
\Delta y_i &= \Delta \frac{pay_i}{pay} \\
&= \frac{(pay_{it-1} + \Delta pay_i)\, pay_{t-1} - (pay_{t-1} + \Delta pay)\, pay_{i,t-1}}{(pay_{t-1} + \Delta pay)\, pay_{t-1}} \\
&= \frac{(pay_{it-1} + \Delta pay_i)\, pay_{t-1} - (pay_{t-1} + \sum_i \Delta pay_i)\, pay_{i,t-1}}{(pay_{t-1} + \sum_i \Delta pay_i)\, pay_{t-1}} \\
&= \frac{pay_{i,t-1} + \Delta pay_i}{pay_{t-1} + \sum_i \Delta pay_i} - \frac{pay_{i,t-1}}{pay_{t-1}} \\
&= \frac{pay_{i,t-1} + \Delta pay_i}{pay_{t-1} + \sum_i \Delta pay_i} - y_{i,t-1},
\end{aligned} \tag{B.19}$$

which depends only on initial $t-1$ data and estimates of Δpay_i for each industry i.

While we are ultimately interested in Δpay_i, we can recover Δpay_i from estimates of $\Delta \ln(pay_i)$, as

$$\begin{aligned}
\ln(\Delta pay_i) &= \ln\left(\frac{pay_{i,t+1}}{pay_{i,t}}\right) \\
\Rightarrow \exp\left[\Delta \ln(pay_i)\right] &= \frac{pay_{i,t+1}}{pay_{i,t}}
\end{aligned} \tag{B.20}$$

and

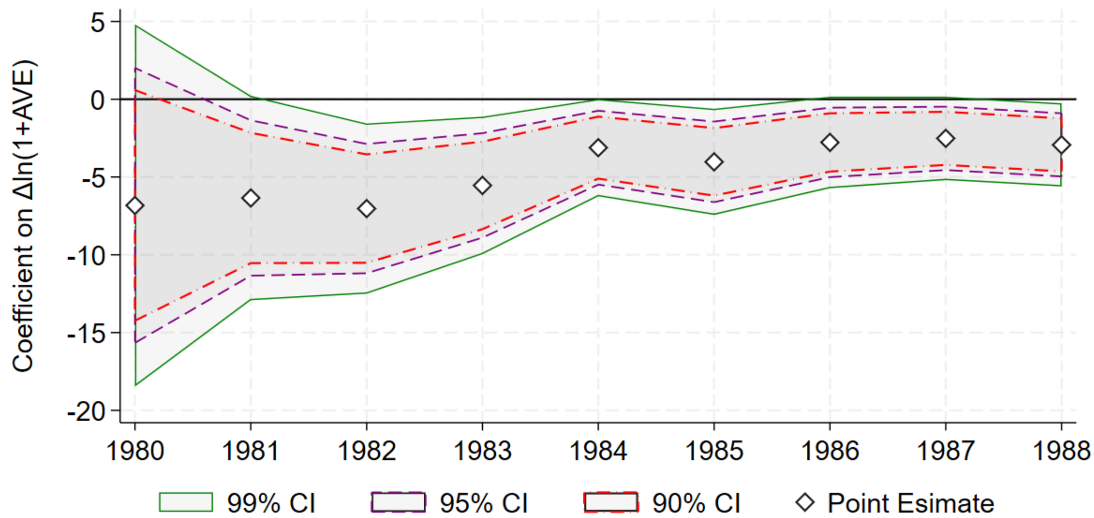
$$
\begin{aligned}
\Delta pay_i &= pay_{i,t+1} - pay_{i,t} \\
&= pay_{i,t} \frac{pay_{i,t+1}}{pay_{i,t}} - pay_{i,t} \\
&= pay_{i,t} \left(\frac{pay_{i,t+1}}{pay_{i,t}} - 1 \right) \\
&= pay_{i,t} \left(\exp \left[\Delta \ln \left(pay_i \right) \right] - 1 \right). \qquad \text{(B.21)}
\end{aligned}
$$

We present the estimated coefficients of our key covariate on both the within industry effects $\Delta y_{NP,i}$ and the between components Δy_i in Figure D.1. As controls we include the full set of controls used throughout the paper.[48]

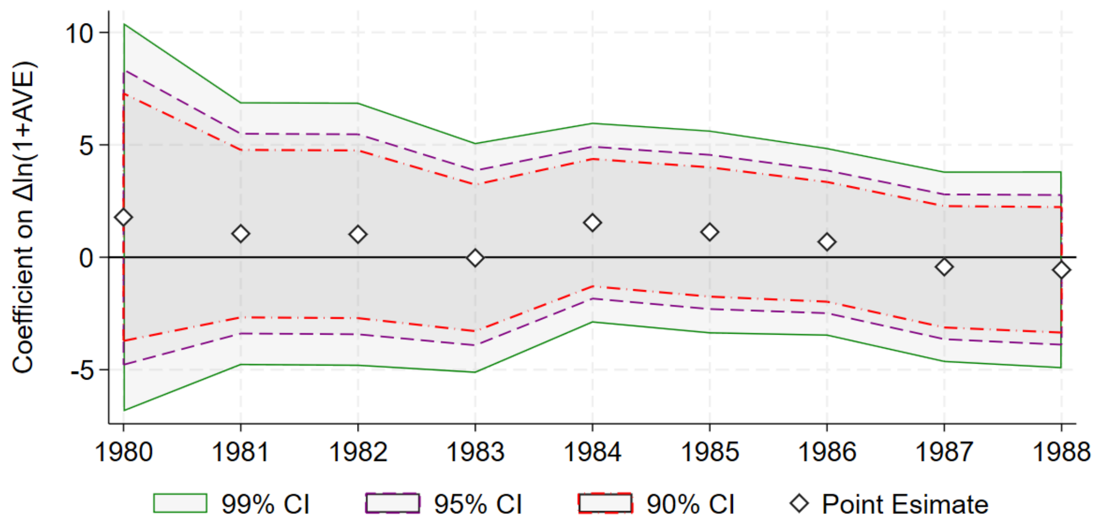
Consistent with the results of this decomposition presented in Figure 7 in the main text, we find no role for the Tokyo round tariffs affecting the aggregate industry share of pay among NBER manufacturing industries.

[48]To facilitate comparison of magnitudes over time, in these specifications we use the observed tariff change between 1978 and each year in constructing the endogenous variable. For the instrument, we allow the eight-year change in AVE predicted by the Swiss Formula to phase in linearly, as was the case in practice.

Figure D.1: Trade Liberalization and Pay Inequality – Between and Within Components



(a) Within Industry Non-Production Pay-Share



(b) Across Industry Pay-Share

Notes: Figure presents IV estimates from estimating equation (1). Each year t point estimate is from a separate specification with the change in the dependent variable defined as the change between 1979 and year t. Dependent variable is within-industry non-production pay share ($y_{NP,i,t}$) in panel (a) industry pay share ($y_{i,t}$) in panel (b). Year t AVE tariff variable and instrumented AVE tariff variable reflect linear phase in of Tokyo Round tariff cuts, as happened in practice. Data measured at or concorded to the 4-digit SIC87 industry classification. Specifications use full controls as in column 4 of Table 3. Robust standard errors. Data from Becker et al. (2021) and Greenland et al. (2025) .