Tariff Pass-Through to U.S. Producer Prices: Firm-Level Evidence From the 2018–19 Trade War*

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Abstract

This paper estimates the pass-through from import tariffs to U.S. producer prices using confidential firm-level data from the Longitudinal Firm Trade Transactions Database (LFTTD) linked with Producer Price Index (PPI) microdata. Exploiting the large and sudden tariff increases during the 2018–19 U.S.–China trade war, we construct firm-specific measures of effective input tariff exposure using each firm's own product-country import portfolio. We address the endogeneity of contemporaneous import weights—which reflect firms' endogenous sourcing responses to tariffs—by constructing Bartik-style instrumental variables that hold import composition fixed at its pre-trade-war (2017) level. Our IV estimates imply that a 1 percentage point increase in effective input tariffs raises output prices by approximately 0.5–0.6 percentage points at a 12-month horizon. To understand the pass-through, we develop a structural model of endogenous supplier choice in which firms optimally adjust via both intensive and extensive margins. We use the estimated pass-through and the structural model to quantify the aggregate inflationary impact of the 2018–19 China tariffs and to evaluate a counterfactual universal tariff motivated by the broad-based tariff policies of 2025–26..

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Keywords: tariffs, pass-through, producer prices, trade war, intermediate inputs, LFTTD, Bartik instrument

*The views expressed in this paper are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Federal Reserve Bank of Atlanta, the Federal Reserve Bank of Boston, or the Federal Reserve System. All remaining errors are our own.

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1 Introduction

A central question in international trade policy concerns the extent to which import tariffs affect domestic prices. The 2018–2019 trade war between the United States and China generated some of the largest tariff increases in modern U.S. history, providing a powerful natural experiment for studying this question. A rapidly growing literature, beginning with Amiti, Redding, and Weinstein (2019), has documented that the tariffs were almost entirely borne by U.S. importers—that is, pass-through to import prices was near-complete. But import prices are only the first link in the domestic price chain. An equally important yet less-studied question is how tariffs on imported intermediate inputs propagate to the prices that domestic producers charge for their output.

This paper provides new firm-level evidence on this question. We combine two confidential Census Bureau datasets: the Longitudinal Firm Trade Transactions Database (LFTTD), which records the universe of U.S. import and export transactions at the firm-HTS10-country-month level, and producer price microdata from the BLS, which tracks the prices that individual manufacturing firms charge for specific products over time. This linkage enables us to observe, for each firm in each month, both its complete portfolio of import transactions—including tariff rates, product codes, and origin countries—and the prices it charges for its domestic output. To our knowledge, this is the first paper to estimate this input-tariff-to-output-price pass-through at the firm level using linked trade-and-price microdata.

Our empirical approach extends the Amiti, Redding, and Weinstein (2019) methodology in two important directions. First, whereas Amiti, Redding, and Weinstein measure tariff exposure at the industry level using publicly available trade data and BEA input-output tables, we exploit the granularity of the LFTTD to construct *firm-level* measures of effective input tariff exposure. For each firm in each month, we compute the effective tariff as an import-weighted average of product-country tariff rates, using the firm's own rolling 12-month import values as weights. Two firms in the same NAICS industry may import very different bundles of inputs from different countries and thus face very different effective tariff rates. Second, we construct industry-level measures of imported input intensity and import share by merging the Census of Manufactures (CMF), the Longitudinal Business Database (LBD), and the LFTTD.

A key identification challenge is that the effective tariff rate constructed from contemporaneous import weights is endogenous. Firms facing higher tariffs on imports from China may respond by sourcing the same intermediate inputs from alternative countries, by replacing imported intermediates with domestically produced substitutes, or by shifting their import mix toward product categories that face lower tariff rates—all of which alter the import weights used to compute the measured effective tariff rate. To address this endogeneity, we construct Bartik-style instrumental variables that hold each firm's import composition fixed at its pre-trade-war (2017) level while allowing tariff rates to vary over time. The "share" component is the firm's 2017 import portfolio—determined before the trade war was announced and therefore free of endogenous sourcing responses—and the "shift" component is the change in de jure (officially announced statutory) tariff rates at the product-country level.

Our main findings are as follows. First, we find positive and statistically significant pass-through from input tariffs to producer prices across all specifications. In our preferred IV estimates using the 2017–2019 baseline sample, a 1 percentage point increase in the effective input tariff rate over a 12-month period raises output prices by approximately 0.49–0.60 percentage points over the same period. Extending the price-change window to 24 months yields estimates of similar magnitude, indicating that the pass-through is largely complete within the first 12-month period and does not continue to build thereafter. Second, the IV estimates are systematically and substantially larger than OLS estimates—0.58 versus 0.35 in the baseline specification—exactly as predicted by our bias analysis. Third, the results are robust to a wide battery of specification checks: weighted versus unweighted regressions, industry-level versus firm-level exposure measures, de jure versus de facto tariff instruments, and the inclusion of industry-by-month fixed effects.

To understand *why* IV exceeds OLS—and to draw policy implications from this gap—we develop a structural model of endogenous supplier choice building on Halpern, Koren, and Szeidl (2015). In the model, each firm imports intermediate inputs from a set of foreign countries, choosing both how much to import from each source (the intensive margin, governed by CES substitution between domestic and foreign varieties) and whether to import from each source at all (the extensive margin, governed by stochastic fixed costs of importing). A tariff increase on country c endogenously reduces the firm's import share from c through both channels. These endogenous choices imply a downward bias in OLS estimates that rely on current import shares

to measure tariff exposures.

We then use the structural model to quantify the aggregate inflationary impact of tariffs. Using the estimated pass-through and the observed distribution of firm-level exposure, we find that the 2018–19 China tariffs raised aggregate manufacturing PPI by approximately 0.2–0.3 percentage points through the input cost channel alone. We also extend the analysis to a counterfactual that is directly motivated by the shift in U.S. trade policy during 2025–2026. The model predicts that a universal tariff of the same rate as the China-only tariff generates an aggregate PPI effect that is 4–5 times larger. The per-unit pass-through is higher, because the substitution "escape valve"—the ability to shift sourcing toward non-tariff countries—is simultaneously shut down across all supplier countries.

This paper contributes to several literatures. We contribute most directly to the literature on the effects of the 2018–2019 trade war (Amiti, Redding, and Weinstein, 2019, 2020; Fajgelbaum et al., 2020; Cavallo et al., 2021; Flaaen, Hortacsu, and Tintelnot, 2020) by tracing the propagation of tariff costs one step further through the production chain—from import prices to domestic output prices. We also contribute to the cost pass-through literature (Goldberg and Knetter, 1997; Gopinath, Itskhoki, and Rigobon, 2010; Amiti, Itskhoki, and Konings, 2014). Finally, our paper draws on the literature that uses Census microdata to study trade. The LFTTD has been used to study trade dynamics (Bernard, Jensen, and Schott, 2009), trade adjustment margins (Monarch, 2021), and distributional effects of trade shocks (Autor, Dorn, and Hanson, 2013). We add to this body of work by linking LFTTD import transactions directly to BLS producer price microdata, a combination that—to the best of our knowledge—has not been exploited before.

The remainder of the paper is organized as follows. Section 2 describes the data sources and the construction of tariff and exposure measures. Section 3 presents the empirical methodology and reports the main results, including robustness checks. Section 4 develops a structural model of endogenous supplier country choice that characterizes how firms adjust both the intensity and the set of their import sources in response to tariff shocks. Section 5 examines the model's implications, using the structural framework to explain the systematic gap between IV and OLS estimates and quantifying the aggregate PPI effect of the China tariffs as well as from a universal tariff counterfactual. Section 6 concludes.

2 Data and Measurement

Our analysis requires linking firms' import transactions to their output prices, which we accomplish by combining four confidential datasets (Section 2.1). A central contribution of this paper is the construction of tariff measures at multiple levels of aggregation: we build product-country tariff rates in both de facto and de jure variants from the LFTTD microdata, aggregate them to industry-level input and output tariff measures following Amiti, Redding, and Weinstein, and construct firm-level effective tariff measures that exploit each firm's own import portfolio (Section 2.3). We then describe the baseline and extended samples used in estimation (Section 2.2).

2.1 Data Sources

We combine four datasets. The *Longitudinal Firm Trade Transactions Database* (LFTTD), maintained by the U.S. Census Bureau, contains transaction-level records from U.S. customs forms, covering the universe of U.S. merchandise imports since 1992. Each import transaction records the customs value and quantity of the shipment, a 10-digit Harmonized Tariff Schedule (HTS10) product code (comprising approximately 17,000 product categories), the country of origin, and identifiers for both the U.S. importing firm and the foreign exporter. Critically for our purposes, the LFTTD also records the duties collected on each transaction, enabling the construction of transaction-level measures of the effective tariff rate as the ratio of duties paid to customs value.

The *Producer Price Index* (PPI) microdata, published by the Bureau of Labor Statistics (BLS), provide transaction-based prices at the individual good level that are representative of the entire U.S. production sector (see, e.g., Nakamura and Steinsson, 2008; Gilchrist, Schoenle, Sim, and Zakrajšek, 2017). The BLS samples goods produced by each firm and uniquely identifies them according to their price-determining characteristics, including the type of buyer, the type of market transaction, the method of shipment, the size and units of shipment, the freight type, and the day of the month of the transaction. This granularity ensures that the PPI captures pure price changes—changes attributable solely to market factors—rather than variation arising from shifts in product characteristics or transaction terms. Following Gilchrist, Schoenle, Sim, and Zakrajšek (2017), we construct the dependent variable as the h-month log (quality-adjusted) price change,

$\Delta_h \log(p_{j,k,i,t}) = \log(p_{j,k,i,t}) - \log(p_{j,k,i,t-h})$.[1] In the paper, we consider both $h = 12$ and $h = 24$. The 12-month change is our baseline specification, and we refer to it simply as $\Delta \log(p_{j,k,i,t})$.

The *Census of Manufactures* (CMF), conducted every five years as part of the U.S. Economic Census, covers the universe of manufacturing establishments with one or more paid employees and provides detailed data on the total value of shipments, annual payroll, production worker wages and hours, cost of materials consumed (by material code), capital expenditures, and energy costs. We merge the CMF with the *Longitudinal Business Database* (LBD) by establishment identifier to aggregate establishment-level records to the firm level and construct measures of firm-level sales, wage bills, and material costs. The 2017 CMF—the most recent vintage available prior to the onset of the 2018–2019 U.S.–China trade war—provides a snapshot of firms' cost structures immediately before the tariff shock.

2.2 Sample

Our baseline sample covers January 2017 to December 2019, the period of active tariff escalation and implementation. The extended sample runs from January 2017 to December 2022, incorporating the partial tariff rollbacks and the COVID-19 pandemic period.

We merge the PPI and LFTTD databases using a multi-step name-matching algorithm designed to maximize match rates while minimizing false positives. Because the two databases do not share a common firm identifier, we rely on business names as the primary matching variable (see Appendix A for more details).

The resulting sample is a panel of approximately 290,000–310,000 firm-product-month observations from roughly 2,900–3,200 unique firms in the baseline sample, and 540,000–580,000 observations from 3,400–3,600 firms in the extended sample. These firms span a broad range of manufacturing industries, including food products,

[1] A distinctive feature of the PPI micro data is the BLS's quality-adjustment procedures. To ensure that the index measures only pure price changes, the BLS excludes any change in the recorded price that is attributable to a change in the "base price," such as physical modifications to the sampled item or changes in the way it is sold. The BLS employs several quality-adjustment methods depending on the type of change and the information available from the respondent (US Bureau of Labor Statistics, 2014). Letting $p^*_{j,k,i,t}$ denote the recorded transaction price of good k produced by firm i and $p^b_{j,k,i,t}$ the corresponding base price, the quality-adjusted good-level price is defined as $p^*_{j,k,i,t} \equiv p^*_{j,k,i,t}/p^b_{j,k,i,t}$ (Gilchrist, Schoenle, Sim, and Zakrajšek, 2017).

chemicals, metals, machinery, electrical equipment, and transportation equipment.

2.3 Tariff Measures

We construct tariff measures at three levels of aggregation. First, we build product-country-month tariff rates in both de facto and de jure variants directly from the LFTTD microdata. These product-country rates are the building blocks for all subsequent measures. Second, we aggregate them to industry-level input and output tariff changes following the methodology of Amiti, Redding, and Weinstein, using BEA input-output tables to trace the propagation of tariffs through the production network. Third, and most novel, we construct firm-level effective tariff measures that exploit each firm's own product-country import portfolio observed in the LFTTD, providing a substantially richer source of cross-sectional variation than the industry-level measures.

Product-Country Tariff Rates

A key innovation of this paper is the construction of firm-level effective and de jure tariff measures. Both rates are constructed from the LFTTD microdata rather than from external tariff schedule databases.

De facto (effective) tariff rates. The LFTTD records, for each import transaction by firm j of HTS10 product h from country c in month t, both the import value $v^{\text{import}}_{j,h,c,t}$ and the duty payment $v^{\text{import}}_{j,h,c,t} \cdot d_{j,h,c,t}$, where $d_{j,h,c,t}$ is the ad valorem tariff rate effectively applied to that shipment. Because different firms importing the same product-country pair may face different effective rates—due to tariff exemptions, duty drawback programs, product exclusions, or differences in rate provision classifications—we aggregate across firms to obtain the product-country-month effective tariff rate:

$$d^{\text{eff}}_{h,c,t} = \frac{\sum_j v^{\text{import}}_{j,h,c,t} \, d_{j,h,c,t}}{\sum_j v^{\text{import}}_{j,h,c,t}}. \tag{1}$$

This import-value-weighted average measures the tariff rate that firms *actually pay* on a given product from a given origin in a given month. It reflects all exemptions, exclusions, and administrative adjustments as they are realized in practice.

De jure (statutory) tariff rates. The de jure tariff rate aims to capture the officially announced statutory rate, abstracting from firm-specific exemptions that reduce the tariff burden below the announced level. To construct this measure, we exploit the *rate provision code* recorded in the LFTTD for each transaction. The rate provision code classifies each import entry according to the specific tariff provision under which it enters the United States—including standard duty rates, various preferential programs, temporary duty suspensions, and tariff exemptions (such as foreign trade zone entries and duty-free provisions).[2] We exclude transactions that enter under exemption provisions and focus on the maximum tariff rate observed across the remaining non-exempt transactions within each (h, c, t) cell. This maximum rate is taken as the de jure tariff rate $d_{h,c,t}^{\text{dejure}}$.

The distinction between de facto and de jure rates is economically important. During the 2018–2019 trade war, the U.S. government announced large tariff increases on Chinese imports, but simultaneously implemented a product exclusion process that exempted certain transactions from the new duties. As a result, the effective tariff rate that firms actually paid could be substantially lower than the announced statutory rate. By constructing both measures from the LFTTD, we can assess whether our pass-through estimates are sensitive to this distinction. As we show below, the de jure and de facto Bartik instruments yield very similar second-stage estimates, indicating that our findings are not driven by the specific treatment of tariff exemptions.

Lastly, we backfill and forward fill missing effective tariff rates to create a balanced panel and compute 12-month changes:

$$\Delta \log\left(1 + d_{h,c,t}^{eff}\right) = \log\left(1 + d_{h,c,t}^{eff}\right) - \log\left(1 + d_{h,c,t-12}^{eff}\right). \tag{2}$$

Similarly, we compute 12-month changes in de jure tariff rates and refer to them as $\Delta \log\left(1 + d_{h,c,t}^{dejure}\right)$.

Output Tariffs

Before we measure input tariffs, we construct an industry-level *output* tariff measure that captures the *competition channel*: the degree to which import tariffs shield domestic producers from foreign competitors selling similar products. When tariffs are imposed on finished goods that compete directly with a domestic industry's output,

[2] See https://www.census.gov/foreign-trade/reference/codes/rp.html.

they raise the effective price of foreign substitutes, reducing competitive pressure and allowing domestic firms to raise their own prices—even absent any change in input costs. To measure this channel, we follow Amiti, Redding, and Weinstein and compute, for each 6-digit NAICS industry i, the import-weighted average change in tariff rates on products classified within that industry:

$$\Delta \log\left(1 + d_{i,t}^{\text{output, eff}}\right) \equiv \sum_{(h,c),\, h \in i} \frac{v_{h,c,0}^{\text{import}}}{\sum_{(h',c'),\, h' \in i} v_{h',c',0}^{\text{import}}} \Delta \log\left(1 + d_{h,c,t}^{eff}\right), \tag{3}$$

where $v_{h,c,0}^{\text{import}} \equiv \sum_j v_{j,h,c,0}^{\text{import}}$ denotes total 2017 U.S. imports of HS product h from country c, and the summation in the denominator aggregates over all product–country pairs (h', c') with h' classified within industry i. The weights are held fixed at their 2017 (pre-trade-war) values to avoid endogeneity from contemporaneous import adjustments. We map HS codes to 6-digit NAICS industries using the Census Bureau's official concordance files.[3]

The output tariff $\Delta \log(1 + d_{i,t}^{\text{output, eff}})$ varies only at the industry-by-month level, in contrast to the firm-level input tariff. An increase in this measure indicates that industry i's domestic products face less foreign competition because tariffs have raised the landed cost of competing imports. In our regression framework, a positive coefficient on the output tariff captures this competition effect: domestic producers exploit the tariff-induced reduction in import competition by raising markups and, consequently, output prices.

Input Tariffs

Import tariffs on intermediate inputs—raw materials, components, and semi-finished goods that domestic manufacturers purchase from abroad—raise the cost of production for firms that rely on imported intermediates. Unlike output tariffs, which affect competitive pressure from foreign rivals, input tariffs operate through the *cost channel*: they increase firms' marginal costs and, to the extent that these cost increases are passed through to customers, raise domestic producer prices. Measuring the input tariff burden that each firm faces is therefore central to estimating the inflationary consequences of trade policy.

[3] Available at https://www.census.gov/foreign-trade/reference/codes/index.html; we use the "HTS_Import_Export_Concordance" file.

A key innovation of this paper is the construction of *firm-level* effective input tariff measures that exploit the granularity of the LFTTD. Because the LFTTD records the universe of U.S. import transactions at the firm-product-country-month level, we can observe each firm's complete import portfolio and compute a tariff exposure measure that reflects the specific products each firm imports, the countries from which it sources, and the relative importance of each product–country pair in the firm's total import bill.

Using the LFTTD, we first compute the effective (de facto) tariff rate at the product-country-month level as the import-value-weighted average across firms, defined in equation (1). For firm j in month t, we measure the rolling 12-month import value $v^{(12m)}_{(h,c)\to j,t}$ of good h from country c between months $t-11$ and t, and compute the effective input tariff level as:

$$\sum_{(h,c)} \frac{v^{(12m)}_{(h,c)\to j,t}}{\sum_{(h,c)} v^{(12m)}_{(h,c)\to j,t}} \log\left(1 + d^{eff}_{h,c,t}\right).$$

Constructing the analogous quantity at $t-12$ using import values between $t-23$ and $t-12$, the 12-month change in the firm's effective input tariff is:

$$\begin{aligned}\Delta \log\left(1 + d^{\text{input,eff}}_{j,t}\right) = &\sum_{(h,c)} \frac{v^{(12m)}_{(h,c)\to j,t}}{\sum_{(h,c)} v^{(12m)}_{(h,c)\to j,t}} \log\left(1 + d^{eff}_{h,c,t}\right) \\ &- \sum_{(h,c)} \frac{v^{(12m)}_{(h,c)\to j,t-12}}{\sum_{(h,c)} v^{(12m)}_{(h,c)\to j,t-12}} \log\left(1 + d^{eff}_{h,c,t-12}\right).\end{aligned} \tag{4}$$

The rolling 12-month window smooths out seasonal fluctuations in trade patterns while remaining responsive to medium-term shifts in sourcing. Two firms classified in the same NAICS industry may import very different bundles of inputs from different countries and thus face very different effective tariff rates—variation that is invisible in industry-level measures.

It is instructive to compare our firm-level input tariff measures with the industry-level approach pioneered by Amiti, Redding, and Weinstein (2019). Because Amiti, Redding, and Weinstein work with publicly available data, they cannot observe individual firms' import portfolios. Instead, they rely on the BEA's 2007 input-output (I-O) table to approximate each industry's exposure to tariffs on imported interme-

diates. Let $v_{i \to i^o}^{\text{input}}$ denote the value of intermediate inputs that downstream industry i^o sources from upstream industry i. The ARW input tariff rate change for industry i^o is:

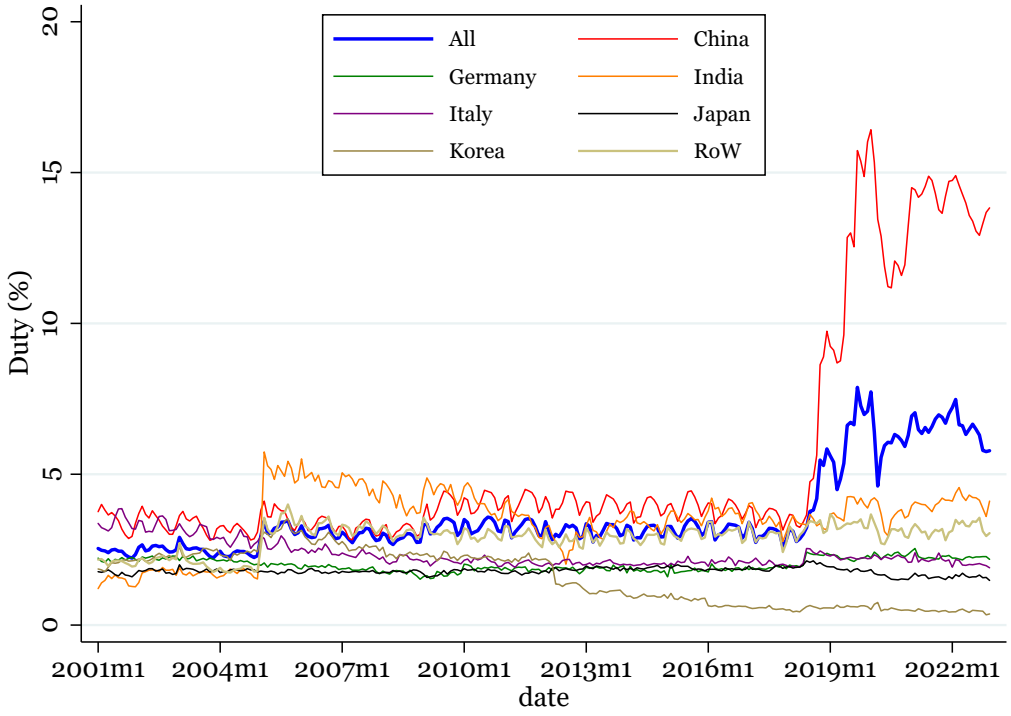
$$\Delta \log \left(1 + d_{i^o,t}^{\text{input, ARW}}\right) \equiv \sum_i \frac{v_{i \to i^o}^{\text{input}}}{v_{i^o}^{\text{input}}} \Delta \log \left(1 + d_{i,t}^{\text{output}}\right), \tag{5}$$

where $v_{i^o}^{\text{input}} \equiv \sum_i v_{i \to i^o}^{\text{input}}$ is total intermediate inputs sourced by industry i^o from all upstream industries, and $\Delta \log(1 + d_{i,t}^{\text{output}})$ is the change in the output tariff for upstream industry i. Intuitively, this measure assigns each downstream industry a weighted average of its upstream industries' output tariffs, with weights proportional to the I-O table's expenditure shares.

Our firm-level measures offer three advantages over this industry-level approach. First, the I-O table captures *average* inter-industry linkages but cannot reflect the substantial within-industry heterogeneity in sourcing patterns. Two firms in the same 6-digit NAICS industry may source intermediates from entirely different countries—one primarily from China, the other from Germany—and therefore face very different tariff exposures during the trade war. Our firm-level measures capture this variation directly. Second, the I-O table does not distinguish between domestically produced and imported intermediates within each upstream industry, nor does it identify the country of origin of imports. Our measures, built from actual firm-level import transactions, incorporate both the product composition and the geographic origin of each firm's imports. Third, the BEA I-O table is updated infrequently (the version used by Amiti, Redding, and Weinstein dates to 2007), so the industry-level weights may not accurately reflect the production network at the time of the trade war. Our firm-level weights are constructed from contemporaneous or near-contemporaneous LFTTD transactions, ensuring that the measured tariff exposure reflects actual sourcing patterns during the relevant period.

Figure 1 plots the value-weighted average ad valorem duty rate on U.S. merchandise imports from 2001 to 2022, broken out by major source countries. From 2001 through early 2018, tariff rates were low and remarkably stable across all source countries, generally hovering below 5 percent. As successive rounds of Section 301 tariffs took effect, tariff rates on imports from China surged from roughly 3 percent to approximately 15–20 percent by late 2019, although no other major trading partner experienced a comparable increase.

Figure 1: Ad Valorem Duty Rates by Source Country



NOTE: This figure plots the value-weighted average ad valorem import duty rate (duties collected as a percentage of customs value) for U.S. merchandise imports from 2001 through 2022, disaggregated by major source country: China, Germany, India, Italy, Japan, Korea, and the rest of the world (RoW). The "All" line shows the simple average across all source countries. The sharp increase in the China series beginning in mid-2018 reflects the successive rounds of Section 301 tariffs imposed during the U.S.–China trade war.
SOURCE: Authors' calculations using the U.S. Census Bureau's Longitudinal Foreign Trade Transactions Database.

Tariff Exposure

To measure the degree to which each industry relies on imported intermediates, we construct two exposure variables: the import share (IS_i) and the imported input intensity (III_i). Both are computed at the industry level using 2017 data from the LFTTD and CMF.

The import share captures the fraction of domestic absorption accounted for by imports:

$$IS_i = \frac{\text{mvalue}_i}{\text{mvalue}_i + \text{sales}_i - \text{xvalue}_i}, \tag{6}$$

where mvalue_i and xvalue_i denote total import and export values for industry i in 2017 from the LFTTD, and sales_i denotes total shipments from the CMF. The denominator is industry-level absorption: domestic shipments net of exports plus imports.

The imported input intensity scales the import share by the share of total variable cost attributable to materials:

$$III_i = \frac{IS_i \times \text{materialcost}_i}{\text{materialcost}_i + \text{wagebill}_i}, \tag{7}$$

where materialcost_i and wagebill_i are total material costs and total wages for industry i from the CMF. The denominator approximates total variable cost (materials plus labor), so III_i measures the fraction of variable cost that comes from imported intermediates.

Summary Statistics. Table 1 reports summary statistics for key variables used in our empirical analysis. From the table, 12-month changes in both input and output tariffs are shown to be right-skewed: the median firm experiences zero or near-zero tariff change, while the 75th percentile shows meaningful positive changes of 0.45 and 0.53 percentage points, respectively, reflecting the concentrated nature of the trade war shock. The table documents substantial variation in import exposure across industries. The mean industry-level imported input intensity is 20.6 percent, indicating that roughly one-fifth of the average manufacturing industry's material costs are sourced from abroad. The interquartile ranges are wide, confirming that import exposure varies substantially across industries. For the interaction terms in the last two rows, their medians are at or near zero, consistent with the fact that many observations experience negligible tariff-exposure products, while the right tail captures firms in import-intensive industries hit by the tariff shock.

3 Empirical Methodology and Results

This section presents our empirical framework and main results. We organize the analysis in four parts. Section 3.1 introduces our baseline specifications—which estimate the unconditional average pass-through from input tariffs to output prices—and reports OLS and IV results. Section 3.2 extends the analysis to exposure-interaction specifications that allow the tariff effect to scale with each industry's reliance on im-

Table 1: Selected Summary Statistics

	Mean	StdDev	P25	P50	P75
Δ_{12} Input tariffs	0.46	3.29	−0.15	0.0	0.45
Δ_{12} Output tariffs	0.54	1.79	−0.02	0.01	0.53
Δ_{12} log PPI (industry)	3.66	11.20	0.0	0.0	5.11
Import share, ind. (IS_i)	30.0	25.2	8.4	23.6	45.0
Imp. input intensity, ind. (III_i)	20.6	16.9	6.6	17.3	31.5
$III_i \times \Delta_{12}$ input tariff	0.11	0.85	−0.02	0.0	0.07
$IS_i \times \Delta_{12}$ output tariff	0.17	0.64	0.00	0.0	0.09

NOTE: This table reports summary statistics for selected variables in the merged LFTTD–PPI panel dataset during 2017–2022. It summarize the regression variables used in Sections 3.1 and 3.2; percentile columns report P25, P50, and P75. Tariff changes are 12-month log changes, expressed in percentage points. Industry-level measures of import share (IS_i) and imported input intensity (III_i), expressed in percent, are constructed from the 2017 CMF, LBD, and LFTTD.

ported intermediates. Section 3.3 discusses the distinct economic interpretations of the two sets of results. Section 3.4 presents robustness checks.

3.1 Baseline Pass-Through Estimates

Our baseline regressions relate 12-month changes in firm-level output prices to changes in effective input tariffs. We consider two specifications that vary in whether the output tariff competition channel is included or absorbed by fixed effects.

Our first specification estimates the pass-through of both input and output tariffs to producer prices:

$$(I): \quad \Delta \log\left(p_{j,k,i,t}\right) = \beta\, \Delta \log\left(1 + d_{j,t}^{\text{input,eff}}\right) + \gamma\, \Delta \log\left(1 + d_{i,t}^{\text{output,eff}}\right) + \eta_j + \delta_t + \epsilon_{j,k,i,t},$$

where $\Delta \log(p_{j,k,i,t})$ is the 12-month log price change for product k of firm j in industry i in month t; $\Delta \log(1 + d_{j,t}^{\text{input,eff}})$ is the 12-month change in the firm's effective input tariff; $\Delta \log(1 + d_{i,t}^{\text{output,eff}})$ is the change in the industry-level output tariff; η_j denotes firm fixed effects; and δ_t denotes time fixed effects. The dependent and independent variables are winsorized at the 0.1% and 99.9% levels. The coefficient β captures the input cost channel (how much of a tariff increase on imported intermediates is passed through to the firm's output price), while γ captures the competition channel (how much the firm raises its output price when tariffs shield it from foreign competition).

The firm fixed effects η_j absorb time-invariant firm characteristics that influence

price-setting behavior—including the firm's average inflation rate, its product mix, its market position, and other persistent features of its cost and demand environment. Since many firms in our sample produce multiple products, the firm fixed effect also controls for the fact that a given firm's prices for different products k may share a common level or trend. Identification of β therefore comes from *within-firm* variation in tariff exposure over time: how the same firm's output prices change as the tariffs on its imported inputs change. Our second specification replaces the time fixed effects and the output tariff variable with industry-by-month fixed effects:

$$(II): \quad \Delta \log\left(p_{j,k,i,t}\right) = \beta\, \Delta \log\left(1 + d_{j,t}^{\text{input,eff}}\right) + \eta_j + \delta_{i,t} + \epsilon_{j,k,i,t},$$

where $\delta_{i,t}$ denotes industry-by-month fixed effects. This is a substantially more demanding specification. The industry-by-month fixed effects absorb *all* time-varying industry-level shocks—including demand shifts, technology changes, sectoral price trends, and, importantly, the output tariff itself (which varies only at the industry-by-month level). Identification of β in this specification comes entirely from *within-industry* variation in firm-level tariff exposure: how firms within the same industry that face different input tariffs (because they import different products from different countries) adjust their output prices differently. A positive and significant β in Specification (II) therefore rules out the possibility that our results are driven by industry-wide confounders correlated with tariff changes.

For each specification, we estimate both unweighted and weighted regressions. The weighted regressions use the BLS importance weight assigned to each price quote in the Producer Price Index program. This weight reflects the item's share in the total value of shipments within its product category, so that price quotes from economically larger product lines receive proportionally more influence in the estimation. The unweighted regressions treat each firm-product-month observation equally, giving equal voice to small and large product lines, and therefore estimate the pass-through for the *typical* firm-product observation in the sample. The weighted regressions, by contrast, estimate the pass-through for the *economically representative* firm-product observation—the one that matters most for the aggregate PPI. If pass-through differs systematically by firm size or product importance (for example, because larger firms have more diversified sourcing or greater market power to absorb cost shocks into margins), the weighted and unweighted estimates will diverge. A finding that

the weighted estimates are modestly smaller than the unweighted estimates would suggest that economically larger product lines exhibit somewhat lower pass-through, consistent with the prediction that firms with larger market shares charge variable markups and absorb a greater fraction of cost shocks (Atkeson and Burstein, 2008; Gopinath, Itskhoki, and Rigobon, 2010). Standard errors are clustered at the firm level throughout to account for the correlation of pricing decisions across products within the same firm and over time.

OLS Results. In specifications (I)–(II), the coefficient β on $\Delta \log(1 + d_{j,t}^{\text{input,eff}})$ measures the *unconditional average* pass-through: by how many percentage points the output price rises when a firm's effective input tariff increases by 1 percentage point, averaging across all importing firms in the sample regardless of their degree of reliance on imported inputs.

Table 2: OLS Regression Results

	Specification (I)		Specification (II)	
	Unwt.	Wt.	Unwt.	Wt.
$\Delta \log(1 + d^{\text{input,eff}})$	0.135	0.131	0.100	0.090
	(0.026)	(0.025)	(0.023)	(0.024)
$\Delta \log(1 + d^{\text{output,eff}})$	0.264	0.287		
	(0.068)	(0.073)		
R^2	0.152	0.177	0.211	0.260
Observations	299k	299k	319k	319k
Firms	3,100	3,100	3,300	3,300
Firm FE	Yes	Yes	Yes	Yes
Industry×Month FE	No	No	Yes	Yes

Notes: Standard errors clustered at the firm level in parentheses. Observation counts are truncated for disclosure avoidance. The dependent variable is the 12-month log change in producer price.

Table 2 presents OLS estimates for the 12-month horizon using the 2017–2019 sample. The coefficient $\hat{\beta}$ on the input tariff variable is positive and statistically significant across all specifications. In specification (I) with firm and time fixed effects, the point estimate is 0.135 (s.e. 0.026) unweighted and 0.131 (0.025) weighted. These estimates imply that, on average, a 10 percentage point increase in a firm's effective

input tariff raises its output price by roughly 1.3 percent over the following 12 months. Moving to specification (II), which replaces time fixed effects with industry-by-month fixed effects $\delta_{i\times t}$, the estimates decline modestly to 0.100 (0.023) and 0.090 (0.024). The attenuation is expected: industry-by-month fixed effects absorb all time-varying industry-level shocks, including the output tariff channel, so that identification comes entirely from within-industry, across-firm variation in tariff exposure. The fact that $\hat{\beta}$ remains positive and highly significant in this demanding specification indicates that the result is not driven by industry-level confounders such as sector-wide demand shifts or common cost trends.

The output tariff coefficient $\hat{\gamma}$ in specification (I) is 0.264 (0.068) unweighted and 0.287 (0.073) weighted, providing evidence for a competition channel: tariffs on goods that directly compete with a firm's products also raise domestic producer prices, as import protection reduces competitive pressure from foreign suppliers. This coefficient is not identified in specification (II), where it is absorbed by the industry-by-month fixed effects.

Endogeneity of Import Weights and IV Strategy

While the OLS estimates provide a useful benchmark, there are reasons to believe they understate the true causal pass-through. A key identification challenge is that the effective tariff rate $\Delta \log(1 + d_{j,t}^{\text{input,eff}})$ is constructed from contemporaneous import weights, which reflect the firm's endogenous sourcing decisions. As we formalize in Section 4, firms facing higher tariffs on imports from a given country may substitute toward alternative suppliers, shift to domestic inputs, or adjust the product composition of their imports—all of which alter the measured effective tariff rate. Because these sourcing adjustments are themselves responses to tariff changes, OLS estimation of Specifications (I)–(II) produces downward-biased estimates of the causal pass-through: firms with the largest tariff increases are precisely those that substitute most aggressively, compressing the measured tariff change and attenuating $\hat{\beta}$ toward zero.

To address this endogeneity, we construct Bartik-style instrumental variables that hold each firm's import composition fixed at its pre-trade-war (2017) level while allowing product-country tariff rates to vary over time. The "share" component is firm j's 2017 import portfolio—determined before the trade war was announced—and the "shift" component is the subsequent change in product-country tariff rates.

Our instrument uses de jure tariff rates as the shift component:

$$\Delta \log\left(1 + d_{j,t}^{\text{IV,dejure}}\right) \equiv \sum_{(h,c)} \frac{v_{(h,c)\to j}^{(2017)}}{\sum_{(h,c)} v_{(h,c)\to j}^{(2017)}} \Delta \log\left(1 + d_{h,c,t}^{\text{dejure}}\right). \tag{8}$$

This instrument has two sources of variation: cross-sectional variation arising from differences in firms' pre-trade-war import portfolios (which products they import, and from which countries), and time-series variation arising from the staggered implementation of tariff increases across product categories and trade-war lists. We favor the de jure instrument because the statutory tariff rates are directly announced by trade policy authorities and are therefore free of the firm-specific exemptions, exclusions, and administrative adjustments that affect de facto rates. The de jure rates thus provide a cleaner measure of the exogenous policy shock.

In the first stage, we regress the endogenous regressor on the corresponding IV, controlling for the same fixed effects as in the second stage. For the exposure-interaction specifications, the first stage regresses $III_i \times \Delta \log(1 + d_{j,t}^{\text{input,eff}})$ on $III_i \times \Delta \log(1 + d_{j,t}^{\text{IV,dejure}})$, along with firm fixed effects and either time or industry-by-time fixed effects.

Identification assumptions. The validity of the Bartik instrument requires two conditions: relevance and exogeneity. Regarding the first condition of relevance, the pre-determined tariff exposure must be a strong predictor of the contemporaneous effective tariff. This is confirmed by the first-stage F-statistics, which uniformly exceed 200 for the de jure IV and 400 for the de facto IV (see Tables 3 and B-1), far above the Stock and Yogo (2005) critical values for weak instruments.

Regarding the second condition of exogeneity, the 2017 import shares must be uncorrelated with subsequent shocks to producer prices, conditional on firm and time fixed effects. This condition is plausible for two reasons. On the "share" side, the import composition was determined before the trade war tariffs were announced (the first Section 301 tariff list was not published until June 2018), so the 2017 import portfolios could not have anticipated the tariff shock. On the "shift" side, the tariff changes are set at the broad product-country level by trade policy authorities in response to geopolitical and macroeconomic considerations, not in response to individual firm pricing decisions. This identification argument follows the framework of **?**,

Table 3: IV Regression Results, De Jure Tariffs

Panel A: Second-stage results	Specification (I)		Specification (II)	
	Unwt.	Wt.	Unwt.	Wt.
$\Delta \log(1 + d^{\text{input,eff}})$	0.203	0.197	0.144	0.121
	(0.035)	(0.036)	(0.033)	(0.036)
$\Delta \log(1 + d^{\text{output,eff}})$	0.244	0.268		
	(0.069)	(0.074)		
Observations	296k	296k	316k	316k
Firms	3,000	3,000	3,200	3,200
Panel B: First-stage results				
$\Delta \log(1 + d^{\text{IV,dejure}})$	0.738	0.740	0.725	0.728
	(0.028)	(0.027)	(0.026)	(0.025)
R^2	0.661	0.659	0.717	0.718
F-statistic	694.1	755.7	806.5	874.0
Firm FE	Yes	Yes	Yes	Yes
Industry×Month FE	No	No	Yes	Yes

Notes: Panel A reports 2SLS estimates using the de jure Bartik IV from equation (8). Panel B reports first-stage results. Standard errors clustered at the firm level. F-statistic is the Kleibergen-Paap weak-instrument test.

who show that Bartik instruments are valid when the shifts (here, tariff changes) are as-good-as-randomly assigned, even if the shares (here, firm-level import portfolios) are endogenous in levels.

IV Results. Table 3 reports 2SLS estimates using the de jure Bartik IV. The IV estimates are systematically larger than OLS, as predicted by the bias analysis of Section 5.1. In specification (I), the IV estimate rises to 0.203 (0.035) unweighted and 0.197 (0.036) weighted—roughly 50 percent larger than the corresponding OLS estimates. In specification (II), the IV estimates are 0.144 (0.033) and 0.121 (0.036). The IV-to-OLS ratio is approximately 1.5 across specifications, consistent with the prediction that endogenous import substitution attenuates OLS toward zero.

The first-stage results (Panel B) confirm a powerful relationship between the Bartik instrument and the endogenous tariff measure. The first-stage coefficient on

$\Delta \log(1 + d^{\text{IV,dejure}})$ is approximately 0.73–0.74, with R^2 values of 0.66–0.72 and F-statistics of 694–874, far exceeding conventional thresholds for instrument relevance. That the first-stage coefficient is below unity is itself informative: it implies that the contemporaneous effective tariff moves less than one-for-one with the pre-determined Bartik instrument, exactly as predicted by endogenous import substitution.

3.2 Pass-Through with Import Exposure

We further refine the analysis by interacting the tariff variables with measures of import exposure. Following Amiti and Heise (2025), we construct industry-level exposure measures from the CMF, LBD, and LFTTD. We merge the CMF and LBD by establishment identifier and retain observations with positive sales in 2017. From the CMF we obtain a firm's wage bills as the sum of salaries and wages (sw) and labor compensation (lc), as well as total material cost (cm). We convert them from thousands of dollars to dollars and refer to them as wagebill_j and materialcost_j, respectively, for firm j. Negative values are replaced by zero.

The exposure interactions allow the tariff effect to scale with each industry's reliance on imported intermediates, separating the *exposure margin* (how much of the firm's cost base is affected) from the *pass-through margin* (how much of the cost increase is reflected in output prices). The exposure-interaction counterparts of Specifications (I) and (II) are:

$$(III): \quad \Delta \log(p_{j,k,i,t}) = \beta \cdot III_i \times \Delta \log\left(1 + d_{j,t}^{\text{input,eff}}\right) + \gamma \cdot IS_i \times \Delta \log\left(1 + d_{i,t}^{\text{output,eff}}\right) + \eta_j + \delta_t + \epsilon_{j,k,i,t},$$

and

$$(IV): \quad \Delta \log(p_{j,k,i,t}) = \beta \cdot III_i \times \Delta \log\left(1 + d_{j,t}^{\text{input,eff}}\right) + \eta_j + \delta_{i,t} + \epsilon_{j,k,i,t}.$$

In Specification (III), the coefficient β on the interaction $III_i \times \Delta \log(1 + d_{j,t}^{\text{input,eff}})$ answers the question: for a firm whose entire material cost base consists of imported intermediates ($III = 1$), what is the output price response to a 1 percentage point tariff increase? The second regressor, $IS_i \times \Delta \log(1 + d_{i,t}^{\text{output,eff}})$, captures the competition channel scaled by the industry's import share. Specification (IV) replaces time fixed effects with industry-by-month fixed effects $\delta_{i,t}$, which absorb all time-

varying industry-level shocks and therefore also absorb the output tariff interaction. As before, all regressions are estimated both unweighted and weighted by the BLS importance weight, and standard errors are clustered at the firm level throughout.

Table 4: OLS with Exposure Interaction

	Specification (III)		Specification (IV)	
	Unwt.	Wt.	Unwt.	Wt.
$III \times \Delta \log(1 + d^{\text{input,eff}})$	0.349	0.313	0.315	0.252
	(0.106)	(0.099)	(0.102)	(0.102)
$IS \times \Delta \log(1 + d^{\text{output,eff}})$	0.691	0.767		
	(0.181)	(0.183)		
R^2	0.151	0.175	0.219	0.262
Obs.	299k	299k	298k	298k
Firms	3,100	3,100	3,100	3,100
Firm FE	Yes	Yes	Yes	Yes
Ind.×Month FE	No	No	Yes	Yes

Notes: Standard errors clustered at the firm level in parentheses. Observation counts truncated for disclosure avoidance. In columns (III)–(IV) the exposure measures III and IS are at the industry level. The dependent variable is the 12-month log change in producer price.

OLS Results. Table 4 presents OLS estimates with the exposure interaction. In specification (III), $\hat{\beta}$ is 0.349 (0.106) unweighted and 0.313 (0.099) weighted. In specification (IV) with industry-by-month fixed effects, the estimates are 0.315 (0.102) and 0.252 (0.102). The output tariff interaction $IS_i \times \Delta \log(1 + d_{i6,t}^{\text{output,eff}})$ is positive and significant in specification (III), with coefficients of 0.691 (0.181) and 0.767 (0.183), indicating that the competition channel also operates more strongly in industries with higher import penetration.

IV Results. Table 5 reports 2SLS estimates using the de jure Bartik IV. The IV estimates are substantially and systematically larger than OLS. In specification (III), $\hat{\beta}$ is 0.577 (0.143) unweighted and 0.531 (0.141) weighted—65 percent larger than the corresponding OLS estimates. With industry-by-month fixed effects in specification (IV), the IV estimates are 0.579 (0.149) and 0.492 (0.160). The stability of the input tariff coefficient across specifications (III) and (IV) is notable: adding the

Table 5: IV with Exposure Interaction, De Jure Tariffs

Panel A: Second-stage results	Specification (III)		Specification (IV)	
	Unwt.	Wt.	Unwt.	Wt.
$III \times \Delta \log(1 + d^{\text{input,eff}})$	0.577	0.531	0.579	0.492
	(0.143)	(0.141)	(0.149)	(0.160)
$IS \times \Delta \log(1 + d^{\text{output,eff}})$	0.612	0.694		
	(0.187)	(0.199)		
Obs.	296k	296k	296k	296k
Firms	3,000	3,000	3,000	3,000
Panel B: First-stage results				
$III \times \Delta \log(1 + d^{\text{IV,dejure}})$	0.669	0.669	0.647	0.644
	(0.040)	(0.041)	(0.042)	(0.043)
R^2	0.671	0.667	0.719	0.718
F-statistic	278.2	264.3	239.9	228.7
Firm FE	Yes	Yes	Yes	Yes
Ind.×Month FE	No	No	Yes	Yes

Notes: Panel A reports 2SLS estimates using the de jure Bartik IV from equation (8). Panel B reports first-stage results. Standard errors clustered at the firm level. F-statistic is the Kleibergen-Paap weak-instrument test.

demanding industry-by-month fixed effects barely changes the point estimate (0.577 versus 0.579), providing strong evidence that the estimated pass-through reflects firm-level variation in tariff exposure rather than industry-level confounders.

The first-stage results (Panel B) show a coefficient of 0.669 on the Bartik IV interaction with F-statistics of 229–278 for industry-level exposure. These large F-statistics rule out weak-instrument concerns.

3.3 Interpreting the Two Sets of Results

The two sets of results—without and with exposure interactions—answer related but distinct economic questions.

The non-interacted specifications (I)–(II) estimate the *unconditional average* pass-through across the population of importing firms. This coefficient blends two margins: the *exposure margin*—how much of the firm's cost base is affected by the tariff—and

the *pass-through margin*—how much of the cost increase is reflected in output prices. The IV estimates from Table 3 are therefore directly informative about the aggregate inflationary impact of a tariff shock: they tell us how much the typical importing manufacturer raises its output price in response to a 1 percentage point increase in its effective input tariff, inclusive of the fact that the typical firm in the sample sources only a fraction of its inputs from abroad.

The interacted specifications (III)–(IV) instead isolate the *structural* pass-through rate conditional on the degree of import exposure, separating the exposure and pass-through margins. The coefficient β on $III \times \Delta \log(1 + d^{\text{input,eff}})$ answers: for a firm whose entire material cost base is imported ($III = 1$), what is the output price response to a 1 percentage point tariff increase? The IV estimates from Table 5 are roughly three times larger than the non-interacted estimates, consistent with the fact that the average III_i in our sample is well below unity. Once the coefficient is expressed per unit of import exposure rather than as an unconditional average, it mechanically scales up.

The difference in magnitudes between the two sets of specifications can be reconciled through the average exposure. Denoting the mean of III_i across the sample as $\overline{III}$, the non-interacted coefficient approximately equals $\hat{\beta}^{\text{no-interaction}} \approx \overline{III} \times \hat{\beta}^{\text{interaction}}$. For the de jure IV, the ratio $0.203/0.577 \approx 0.35$ provides an estimate of the effective average imported input intensity in the sample, consistent with the fact that imported intermediates represent a meaningful but far from dominant share of total material costs for the typical U.S. manufacturer.

This decomposition has a practical policy implication. For assessing the aggregate inflationary impact of a tariff shock across the manufacturing sector, the non-interacted estimates (0.14–0.21 per percentage point of tariff increase) are the directly relevant numbers. For understanding the cost pass-through mechanism and predicting the inflationary impact on *specific* firms or industries with known import intensities, the interacted estimates (0.49–0.60 per unit of III) are more informative, as they can be scaled by the firm's or industry's actual III to obtain a tailored pass-through prediction.

3.4 Robustness

We subject our main findings to a battery of robustness checks along five dimensions: the price-change horizon, the sample period, and the choice of tariff instrument. Throughout, we focus on the exposure-interaction specifications (III)–(IV) with industry-level III and IS. The key finding—positive, significant, and economically meaningful pass-through from input tariffs to output prices, with IV estimates systematically exceeding OLS—is robust across all variations.

Twenty-four-month price-change horizon. Our baseline measures the output price response at a 12-month horizon. Table 6 extends the analysis to a 24-month horizon, in which the dependent variable is the price change between month $t-11$ and month $t+12$, while the right-hand-side tariff variable remains the 12-month change $\Delta_{12}\log(1+d_{j,t}^{\text{input,eff}})$.

The 24-month IV estimates are remarkably similar to their 12-month counterparts. In the de jure IV, specification (III), the 24-month estimate is 0.624 (0.152) compared to 0.577 (0.143) at 12 months; in the de facto IV, it is 0.688 (0.147) compared to 0.600 (0.145). These estimates are statistically indistinguishable from the 12-month results. The 24-month results confirm that the vast majority of tariff pass-through is completed within the first year. The modest increase in point estimates at 24 months is consistent with a small residual adjustment beyond 12 months, but the difference is well within sampling uncertainty.

Two additional patterns deserve note. First, the IV-to-OLS gap persists at the longer horizon: in specification (III), the de jure IV estimate of 0.624 exceeds the OLS estimate of 0.482 by approximately 30%, confirming that endogenous import substitution continues to attenuate OLS at longer horizons. Second, the output tariff coefficient $\hat{\gamma}$ in specification (III) is 0.507 (0.167) under the de jure IV, comparable to the 12-month estimate of 0.612 (0.187), indicating that the competition channel also operates at similar magnitudes over the longer horizon.

Extended sample: 2017–2022. Table 7 extends the sample from 2017–2019 to 2017–2022, incorporating the partial tariff rollbacks under the Phase One trade agreement and the COVID-19 pandemic. This extension serves two purposes: it tests whether the pass-through relationship is stable beyond the initial tariff escalation

Table 6: Regression Results with Exposure Interaction (24-month)

	Specification (III)		Specification (IV)	
	Unwt.	Wt.	Unwt.	Wt.
Panel A: OLS				
$III \times \Delta \log(1 + d^{\text{input,eff}})$	0.482	0.464	0.313	0.270
	(0.107)	(0.112)	(0.103)	(0.109)
$IS \times \Delta \log(1 + d^{\text{output,eff}})$	0.555	0.565		
	(0.168)	(0.178)		
R^2	0.223	0.249	0.285	0.328
Obs.	246k	246k	245k	245k
Firms	2,600	2,600	2,600	2,600
Panel B: IV (de jure)				
$III \times \Delta \log(1 + d^{\text{input,eff}})$	0.624	0.619	0.389	0.326
	(0.152)	(0.172)	(0.163)	(0.182)
$IS \times \Delta \log(1 + d^{\text{output,eff}})$	0.507	0.513		
	(0.167)	(0.177)		
Obs.	244k	244k	243k	243k
Firms	2,600	2,600	2,600	2,600
Panel C: IV (de facto)				
$III \times \Delta \log(1 + d^{\text{input,eff}})$	0.688	0.670	0.424	0.356
	(0.147)	(0.169)	(0.160)	(0.179)
$IS \times \Delta \log(1 + d^{\text{output,eff}})$	0.484	0.496		
	(0.167)	(0.177)		
Obs.	244k	244k	243k	243k
Firms	2,600	2,600	2,600	2,600
Firm FE	Yes	Yes	Yes	Yes
Ind.×Month FE	No	No	Yes	Yes

Notes: See notes to Table 4. Sample period: 2017–2019.

period, and it nearly doubles the sample size (from approximately 296k to 548k observations).

All qualitative patterns are preserved. The IV estimates remain positive and statistically significant, and they continue to exceed OLS systematically. In the de jure IV, specification (III), the estimate is 0.556 (0.154) for 2017–2022 versus 0.577 (0.143) for 2017–2019—a modest attenuation of about 4% that is well within the

Table 7: Regression Results with Exposure Interaction (2017–2022)

	Specification (III)		Specification (IV)	
	Unwt.	Wt.	Unwt.	Wt.
Panel A: OLS				
$III \times \Delta \log(1 + d^{\text{input,eff}})$	0.213	0.220	0.167	0.142
	(0.096)	(0.090)	(0.081)	(0.083)
$IS \times \Delta \log(1 + d^{\text{output,eff}})$	0.359	0.382		
	(0.153)	(0.162)		
R^2	0.209	0.217	0.322	0.357
Obs.	565k	565k	565k	565k
Firms	3,600	3,600	3,600	3,600
Panel B: IV (de jure)				
$III \times \Delta \log(1 + d^{\text{input,eff}})$	0.556	0.510	0.500	0.384
	(0.154)	(0.150)	(0.156)	(0.161)
$IS \times \Delta \log(1 + d^{\text{output,eff}})$	0.255	0.300		
	(0.162)	(0.171)		
Obs.	548k	548k	548k	548k
Firms	3,400	3,400	3,400	3,400
Panel C: IV (de facto)				
$III \times \Delta \log(1 + d^{\text{input,eff}})$	0.489	0.445	0.446	0.352
	(0.142)	(0.140)	(0.149)	(0.154)
$IS \times \Delta \log(1 + d^{\text{output,eff}})$	0.279	0.322		
	(0.161)	(0.170)		
Obs.	548k	548k	548k	548k
Firms	3,400	3,400	3,400	3,400
Firm FE	Yes	Yes	Yes	Yes
Ind.×Month FE	No	No	Yes	Yes

Notes: See notes to Table 4. Sample period: 2017–2022.

confidence intervals of both estimates.

The attenuation is expected for two reasons. First, the post-2019 period saw partial tariff rollbacks and the granting of additional product exclusions, which reduced the cross-sectional variation in tariff exposure that identifies the pass-through coefficient. Second, the COVID-19 pandemic introduced large supply and demand shocks—container shipping disruptions, lockdown-induced demand shifts, commodity

price swings—that added substantial noise to the pricing relationship and may have temporarily altered firms' pricing behavior. That the estimates remain significant and of similar magnitude despite this dilution speaks to the robustness of the underlying pass-through mechanism.

Alternative IV: de facto Tariffs. As a robustness check, we also construct an instrument using de facto (effectively collected) tariff rates, which uses the same 2017 import shares but replaces the statutory tariff rates with the import-value-weighted average rates actually collected at the border. Because the de facto rates reflect product exclusions and other administrative adjustments that were partly determined after the trade war began, this instrument is arguably less clean than the de jure version. We include it primarily to verify that our results are not sensitive to the distinction between announced and collected tariff rates—a concern given the extensive exclusion process that accompanied the Section 301 tariffs. The near-identical second-stage estimates across all specifications, reported in Tables B-1 and B-2 in Appendix B, provide reassuring evidence that our findings are not sensitive to the specific treatment of tariff exceptions.

4 A Model of Endogenous Supplier Country Choice

In this section, we develop a model with endogenous supplier country choice. We build on Halpern, Koren, and Szeidl (2015) to model the cost structure of the firm and its choice to import intermeidate inputs. In the next section, we examine its implications, for example, those on the souces of the OLS bias, among others.

Production technology. Time is discrete and indexed by $t = 0, 1, 2, \ldots$ There are N foreign countries indexed by $c = 1, \ldots, N$. Each firm j produces a differentiated variety using labor $L_{j,t}$ and a composite intermediate input $X_{j,t}$ according to:

$$Y_{j,t} = \Omega_{j,t} \, X_{j,t}^{\phi} \, L_{j,t}^{1-\phi}, \tag{9}$$

where $\Omega_{j,t}$ is firm-specific productivity and $\phi \in (0, 1)$ is the intermediate input share. The intermediate input bundle aggregates N types of inputs, each corresponding to

a foreign country:

$$X_{j,t} = \prod_{c=1}^{N} X_{j,c,t}^{\gamma_c}, \qquad \sum_{c=1}^{N} \gamma_c = 1, \tag{10}$$

where γ_c represents the importance of type-c intermediate input. Each type c comes in domestic and foreign varieties that are imperfect substitutes:

$$X_{j,c,t} = \left[Z_{j,c,t}^{\zeta/(1+\zeta)} + a_c^{1/(1+\zeta)} M_{j,c,t}^{\zeta/(1+\zeta)} \right]^{(1+\zeta)/\zeta}, \tag{11}$$

where $Z_{j,c,t}$ and $M_{j,c,t}$ are the quantities of domestic and imported varieties respectively, $a_c > 0$ is an efficiency parameter governing the productivity advantage of the foreign variety, and $\zeta > 0$ is the elasticity of substitution between domestic and foreign varieties.

The price of the foreign variety c in domestic currency is:

$$P_{c,t}^{M} = e_t \, U_{c,t} \, (1 + d_{c,t}), \tag{12}$$

where e_t is the exchange rate, $U_{c,t}$ is the f.o.b. price in foreign currency, and $d_{c,t} \geq 0$ is the ad valorem tariff rate. We denote $V_{c,t}^{*}$ as the price of the domestic variety of type-c intermediate input.

Proposition 1 (Total Variable Cost)**.** *Given the Cobb-Douglas production function* (9)*, the Cobb-Douglas aggregation of intermediate types* (10)*, and the CES substitution between home and foreign varieties* (11)*, the total variable cost function for a firm that is importing from the set $\mathcal{J}_{j,t}$ is given by*

$$TVC_{j,t}(Y_{j,t} \mid \mathcal{J}_{j,t}) = \frac{C_t^{*}}{B_{j,t}^{\phi}} \cdot \frac{Y_{j,t}}{\Omega_{j,t}}, \tag{13}$$

where C_t^{*} and V_t^{*} are the unit cost index and price index for a non-importing firm:

$$C_t^{*} = \left(\frac{V_t^{*}}{\phi} \right)^{\phi} \left(\frac{W_t}{1-\phi} \right)^{1-\phi}, \tag{14}$$

$$V_t^{*} = \prod_{c=1}^{N} \left(\frac{V_{c,t}^{*}}{\gamma_c} \right)^{\gamma_c}, \tag{15}$$

and $B_{j,t}$ is the cost-reduction factor from importing:

$$B_{j,t} = \prod_{c \in \mathcal{J}_{j,t}} b_{c,t}^{\gamma_c}, \tag{16}$$

and $b_{c,t}$ captures the productivity-enhancing effect of importing type-c intermediates:

$$b_{c,t} = \left[1 + a_c \left(\frac{P_{c,t}^M}{V_{c,t}^*} \right)^{-\zeta} \right]^{1/\zeta}. \tag{17}$$

Proof. See Appendix C.

For a non-importing firm, total variable cost is given by

$$TVC_{j,t}^0 = C_t^* \cdot \frac{Y_{j,t}}{\Omega_{j,t}},$$

and marginal cost is $c_{j,t}^0 \equiv C_t^*/\Omega_{j,t}$. By contrast, for an importing firm with a non-empty importing set $\mathcal{J}_{j,t}$, the total variable cost is reduced by a factor of $B_{j,t}^{\phi}$.

It is worthwhile to point out that the expenditure share of imports in type-c intermediate input is:

$$\alpha_{c,t} = \frac{P_{c,t}^M \cdot M_{j,c,t}}{P_{c,t}^M \cdot M_{j,c,t} + V_{c,t}^* \cdot Z_{j,c,t}} = \frac{a_c \left(P_{c,t}^M / V_{c,t}^* \right)^{-\zeta}}{1 + a_c \left(P_{c,t}^M / V_{c,t}^* \right)^{-\zeta}}. \tag{18}$$

Note that $\alpha_{c,t} \in (0,1)$ and is *decreasing* in the relative import price $P_{c,t}^M/V_{c,t}^*$: as imports become more expensive (e.g., due to a tariff increase), the firm shifts spending toward the domestic variety.

Similarly, the share of country c in firm j's total import spending is given by:

$$s_{j,c,t} \equiv \frac{P_{c,t}^M \, M_{j,c,t}}{\sum_{c'=1}^N P_{c',t}^M \, M_{j,c',t}} = \frac{\gamma_c \, \alpha_{c,t} \, \mathbf{1}_{j,c,t}}{\sum_{c'=1}^N \gamma_{c'} \, \alpha_{c',t} \, \mathbf{1}_{j,c',t}}, \tag{19}$$

where $\mathbf{1}_{j,c,t} \equiv \mathbf{1}[c \in \mathcal{J}_{j,t}]$ is the importing indicator. This expression makes transparent the two channels through which tariffs affect import shares. The within-type import expenditure share $\alpha_{c,t}$ responds to tariffs through the intensive margin (CES substitution between domestic and foreign varieties within type c), while the importing indicator $\mathbf{1}_{j,c,t}$ responds through the extensive margin (the discrete decision to

import from country c at all, governed by the stochastic fixed costs). In particular, $\mathbf{1}_{j,c,t}$ is endogenous to tariff changes, as we derive next.

Endogenous Supplier Choice

Importing from country c requires paying a fixed cost $\varepsilon\, C_{F,t}$ per period, where $C_{F,t} = W_t\, \bar{f}$ is a baseline fixed cost proportional to the wage and ε is an i.i.d. draw from a distribution $F(\varepsilon)$ with $\mathbb{E}(\varepsilon) = 1$. The stochastic fixed cost smooths the discrete importing decision and generates a well-defined expected cost function.

Define the reduced-form cost-reduction parameter:

$$\tilde{b}_{c,t} = b_{c,t}^{\gamma_c \phi}. \tag{20}$$

Sort all supplier countries such that $\tilde{b}_{1,t} \geq \tilde{b}_{2,t} \geq \cdots \geq \tilde{b}_{N,t}$. The total variable cost from importing from the best k suppliers is:

$$TVC_{j,t}^{k} = \frac{C_t^*}{\tilde{B}_{k,t}\, \Omega_{j,t}}\, Y_{j,t}, \tag{21}$$

where $\tilde{B}_{k,t} = \prod_{c=1}^{k} \tilde{b}_{c,t}$. Note that $\tilde{B}_{k,t} > 1$ measures the cumulative cost reduction from importing from the best k countries. To reduce variable costs, firm j has an incentive to import from as many supplier countries as possible. However, it also faces a fixed cost $\varepsilon\, C_{F,t}$ per period for each country from which it imports. The firm therefore chooses to import from exactly k suppliers if and only if the realized fixed cost draw satisfies

$$TVC_{j,t}^{k} - TVC_{j,t}^{k+1} < \varepsilon\, C_{F,t} < TVC_{j,t}^{k-1} - TVC_{j,t}^{k}. \tag{22}$$

The first inequality states that the additional cost reduction from importing from country $k+1$ is insufficient to cover the fixed cost, while the second ensures that the marginal benefit from importing from country k is large enough to justify it.

For $k \in \{1, \cdots, N\}$, define the cutoffs

$$\varepsilon_{j,t}^{k} \equiv \frac{TVC_{j,t}^{k-1} - TVC_{j,t}^{k}}{C_{F,t}} = \frac{(\tilde{b}_{k,t} - 1)\, TVC_{j,t}^{k}}{C_{F,t}}, \tag{23}$$

where the second equality uses the fact that $TVC_{j,t}^{k} = TVC_{j,t}^{k-1} / \tilde{b}_{k,t}$. It is straightfor-

ward to show that the cutoffs are strictly decreasing[4]

$$\varepsilon_{j,t}^1 > \varepsilon_{j,t}^2 > \cdots > \varepsilon_{j,t}^N > 0.$$

Intuitively, because suppliers are ranked by their cost-reduction advantage, each successive supplier offers a smaller marginal variable-cost saving, so only firms drawing a lower fixed cost find it worthwhile to add the k-th source. We summarize the above findings about the firm's endogenous sourcing decision in the proposition below.

Proposition 2 (Endogenous Sourcing)**.** *Firm j imports from the best k supplier countries if and only if the realized fixed cost draw satisfies $\varepsilon_{j,t}^{k+1} < \varepsilon < \varepsilon_{j,t}^k$, where the cutoffs are given in (23) and we denote $\varepsilon_{j,t}^{N+1} = 0$.*

The above proposition characterizes a firm's endogenous importing decision. The key insight is that firms optimally source from a *nested* set of supplier countries ranked by their cost-reduction advantage $\tilde{b}_{c,t}$. A firm with a low fixed cost draw ε_j imports from many countries—it can afford the fixed cost of maintaining multiple supplier relationships and benefits from the associated cost reductions. Conversely, a firm with a high fixed cost draw imports from fewer countries or not at all: only the most cost-advantageous suppliers generate sufficient variable cost savings to justify the fixed cost of entry. Crucially, the set of active suppliers always takes the form of a "top-k list": a firm that finds it profitable to import from the third-ranked country necessarily also imports from the first and second, because those countries offer even larger cost advantages per unit of fixed cost. The cutoff $\varepsilon_{j,t}^k$ represents the maximum fixed cost at which a firm would still find it worthwhile to add the k-th supplier; firms with ε_j below this threshold import from at least k countries. Because tariff changes alter the cost-reduction parameters $\tilde{b}_{c,t}$, they shift these cutoffs and thereby change the extensive margin of importing—the number of countries from which a firm sources.

[4]This is because

$$\frac{\varepsilon_{j,t}^{k+1}}{\varepsilon_{j,t}^k} = \frac{(\tilde{b}_{k+1,t} - 1)\, TVC_{j,t}^{k+1}}{(\tilde{b}_{k,t} - 1)\, TVC_{j,t}^k} = \frac{\tilde{b}_{k+1,t} - 1}{(\tilde{b}_{k,t} - 1)\, \tilde{b}_{k+1,t}} = \frac{1}{\tilde{b}_{k+1,t}} \frac{\tilde{b}_{k+1,t} - 1}{\tilde{b}_{k,t} - 1} < 1,$$

where the inequality holds because $1 < \tilde{b}_{k+1,t} < \tilde{b}_{k,t}$.

Expected Marginal Cost. We now wish to construct the expected total cost function, taking expectations over ϵ. The probability of exactly k suppliers is $F\left(\epsilon_{j,t}^{k}\right) - F\left(\epsilon_{j,t}^{k+1}\right)$, and we can write the expected variable cost and the expected fixed cost as follows

$$\begin{aligned}\mathbb{E}[TVC_{j,t}] &= \left(1 - F\left(\epsilon_{j,t}^{1}\right)\right) TVC_{j,t}^{0} + \sum\nolimits_{k=1}^{N} \left(F\left(\epsilon_{j,t}^{k}\right) - F\left(\epsilon_{j,t}^{k+1}\right)\right) TVC_{j,t}^{k}, \\ \mathbb{E}[TFC_{j,t}] &= \sum\nolimits_{k=1}^{N} \left(F\left(\epsilon_{j,t}^{k}\right) - F\left(\epsilon_{j,t}^{k+1}\right)\right) \mathbb{E}\left(\epsilon |\, \epsilon_{j,t}^{k+1} \leq \epsilon < \epsilon_{j,t}^{k}\right) \left(kC_{F,t}\right).\end{aligned}$$

We can further show that the expected total and marginal cost functions can be expressed as a combination of $TVC_{j,t}^{0}$, the total variable cost of a non-importing firm, minus the cost reduction from importing which can be expressed as a function of the extensive margin cutoffs $\varepsilon_{j,t}^{k}$. This is stated in the following proposition:

Proposition 3 (Expected Marginal Cost)**.** *Given the endogenous supplier choice, the expected total cost and marginal cost are given by*

$$\mathbb{E}[TC_{j,t}] = \mathbb{E}[TVC_{j,t}] + \mathbb{E}[TFC_{j,t}] = TVC_{j,t}^{0} - C_{F,t} \sum\nolimits_{k=1}^{N} G(\varepsilon_{j,t}^{k}), \tag{24}$$

$$\mathbb{E}[MC_{j,t}] = c_{j,t}^{0} - c_{j,t}^{f} \sum\nolimits_{k=1}^{N} \varepsilon_{j,t}^{k} F(\varepsilon_{j,t}^{k}), \tag{25}$$

where $TVC_{j,t}^{0} \equiv \frac{C_t^*}{\Omega_{j,t}} Y_{j,t}$ *denotes the total variable cost for a non-importing firm, and*

$$G(\varepsilon_{j,t}^{k}) = \int_{0}^{\varepsilon_{j,t}^{k}} (\varepsilon_{j,t}^{k} - \varepsilon)\, dF(\varepsilon) \tag{26}$$

captures the expected cost savings from the option to import variety k*, and* $c_{j,t}^{0}$ *and* $c_{j,t}^{f}$ *are defined below:*

$$c_{j,t}^{0} = \frac{TVC_{j,t}^{0}}{Y_{j,t}} = \frac{C_t^*}{\Omega_{j,t}} \text{ and } c_{j,t}^{f} = \frac{C_{F,t}}{Y_{j,t}}. \tag{27}$$

Proof. See Appendix C.

Tariff Shock and Marginal Cost Elasticity

To examine the model's implication about the U.S.-China trade war in 2018–2019, we assume that there is a sharp increase in the tariff rate $d_{c^*,t}$ for country c^*.

As a simplifying assumption, we assume that the ranking of countries by their cost-reduction parameters $\tilde{b}_{1,t} \geq \tilde{b}_{2,t} \geq \cdots \geq \tilde{b}_{N,t}$ does not change in response to the tariff shock. First, it is straightforward to show (see Appendix C):

$$\frac{d\tilde{b}_{c,t}}{dd_{c,t}} = -\frac{\phi\,\gamma_c\,\alpha_{c,t}\,\tilde{b}_{c,t}}{1+d_{c,t}}. \tag{28}$$

The sign is negative: a tariff increase raises $P^M_{c,t}$, reduces the cost advantage of importing ($b_{c,t}$ falls), and therefore reduces $\tilde{b}_{c,t}$. The magnitude depends on $\alpha_{c,t}$—the more the firm spends on imports from country c, the larger the cost impact.

How does a tariff hike in country c affects firms' importing decision? From equation (C-23) in Appendix C, we show that

$$\frac{d\varepsilon^k_{j,t}}{d\tilde{b}_{c,t}} = \begin{cases} 0 & \text{if } k < c, \\ c^0_{j,t}/(c^f_{j,t}\,\tilde{B}_{c-1,t}\,\tilde{b}^2_{c,t}) & \text{if } k = c, \\ -\varepsilon^k_{j,t}/\tilde{b}_{c,t} & \text{if } k > c. \end{cases} \tag{29}$$

The above results have a transparent economic interpretation. For $k < c$, the cutoffs of higher-ranked suppliers are unaffected because their cost-reduction parameters and cumulative cost bases do not involve country c—a tariff on an inferior supplier has no bearing on whether it is profitable to import from a superior one. For $k = c$, the own-country cutoff moves in the same direction as $\tilde{b}_{c,t}$: a tariff increase on country c lowers $\tilde{b}_{c,t}$, which reduces $\varepsilon^c_{j,t}$ and thereby shrinks the set of firms that find it profitable to import from c. This is the intuitive "direct" effect: when importing from a country becomes more expensive, fewer firms choose to do so. The most interesting case is $k > c$. The cutoffs for lower-ranked suppliers move in the *opposite* direction of $\tilde{b}_{c,t}$: a tariff on country c that reduces $\tilde{b}_{c,t}$ causes $\varepsilon^k_{j,t}$ to *rise* for all $k > c$. The mechanism is that a weaker country c worsens the cumulative cost base $\tilde{B}_{k,t}$ for multi-sourcing bundles that include c, which raises the incremental benefit of adding the next supplier on the margin. In other words, when the best supplier deteriorates, the option value of diversifying into additional suppliers increases—the crowding-out effect of Proposition 4 operating in reverse.

Proposition 4 (Marginal Cost Elasticity)**.** *Under uniformly distributed fixed costs (i.e. $F(\varepsilon) = \varepsilon/\bar{\varepsilon}$ for $\varepsilon \in (0, \bar{\varepsilon})$), the elasticity of expected marginal cost with respect*

to the tariff factor on imports from country c is given by:

$$\frac{d\log \mathbb{E}[MC_{j,t}]}{d\log(1+d_{c,t})} = \frac{2\,\phi\,\gamma_c\,\alpha_{c,t}}{\bar{\varepsilon}\,\mathbb{E}[MC_{j,t}]}\left[c^0_{j,t}\frac{\varepsilon^c_{j,t}}{\tilde{B}_{c,t}} - c^f_{j,t}\sum_{k=c+1}^{N}\left(\varepsilon^k_{j,t}\right)^2\right] > 0. \tag{30}$$

Proof. See Appendix C.

Proposition 4 provides a closed-form expression for the key object of interest: the elasticity of a firm's expected marginal cost with respect to a country-specific tariff change. This elasticity is the structural analog of the pass-through coefficient β estimated in Section 3. Several features of equation (30) deserve emphasis. First, the elasticity is not a single parameter but a function of the firm's entire import portfolio—the cost-reduction parameters $\tilde{b}_{c,t}$, the import expenditure shares $\alpha_{c,t}$, and the importing cutoffs $\varepsilon^k_{j,t}$—all of which are endogenous to the tariff. Second, the expression inside the brackets is the difference of two terms: a positive term proportional to $\varepsilon^c_{j,t}$ (reflecting the direct cost increase from a tariff on country c) and a negative term summing over all lower-ranked suppliers $k > c$ (reflecting the offsetting cost savings from the expansion of multi-sourcing, i.e., the crowding-out effect operating in reverse). Under the assumption of a uniform fixed cost distribution, the net sign is positive— the direct effect of the tariff increase on the marginal cost of importing from country c exceeds the savings from adjusting the extensive margins of lower ranked suppliers.[5] Third, the elasticity scales with $\phi\gamma_c\alpha_{c,t}$, so pass-through is larger for firms that spend a greater share of their input budget on imports from the tariffed country, consistent with our exposure-interaction estimates in Section 3.2.

4.1 Intensive vs. Extensive Margin Adjustments

The structural model makes transparent why the effective tariff measure constructed from contemporaneous import weights is endogenous: a tariff increase on country c reduces $s_{j,c,t}$ through two channels of intensive and extensive margins.

Intensive margin. Conditional on continuing to import from country c, the firm substitutes toward the domestic variety within the CES composite (11). The within-

[5] Although a tariff increase raises expected total cost, under alternative distributional assumptions it is possible to obtain a reduction in expected marginal cost in response to a tariff increase.

type import expenditure share falls:

$$\alpha_{c,t} = \frac{a_c \left(P^M_{c,t}/V^*_{c,t}\right)^{-\zeta}}{1 + a_c \left(P^M_{c,t}/V^*_{c,t}\right)^{-\zeta}}, \tag{31}$$

which is decreasing in $P^M_{c,t}$ and therefore decreasing in $d_{c,t}$. A tariff hike on country c causes the firm to shift its spending toward the domestic alternative, reducing the observed import share from country c.

Extensive margin. As shown in (32), the tariff increase also causes some firms to drop country c entirely as a supplier, setting $s_{j,c,t} = 0$. For these firms, the tariff on country c no longer appears in their effective tariff measure at all, even though the tariff shock has genuinely raised their costs.

The key mechanism is that tariffs shift the importing cutoffs. Specifically, a tariff increase on country c raises $P^M_{c,t}$ and therefore reduces $b_{c,t}$ and $\tilde{b}_{c,t}$: the cost advantage of importing from country c shrinks. This in turn lowers the cutoff $\varepsilon_{c,t}$ in (23), meaning that fewer firms find it profitable to import from country c. Formally, the change in the probability of importing from country c in response to a tariff increase is:

$$\frac{dF(\varepsilon_{c,t})}{d\log(1+d_{c,t})} = f(\varepsilon_{c,t}) \cdot \frac{d\varepsilon_{c,t}}{d\log(1+d_{c,t})} < 0. \tag{32}$$

This extensive-margin adjustment—firms dropping country c as a supplier and switching to domestic or third-country alternatives—is the structural analogue of the import substitution that generates endogeneity in the OLS regression.

To see how such bias can arise, we derive the first-order approximation of the structural import share $s_{j,c,t}$ and show that the change decomposes into intensive-margin, extensive-margin, and cross-country rebalancing terms. Specifically, we take a first-order approximation of (19) around the pre-trade-war values which, for simplicity, are assumed to take the steady-state values (e.g., $\alpha_{c,2017} = \bar{\alpha}_c$, $\mathbf{1}_{j,c,2017} = \bar{\mathbf{1}}_{j,c}$, $s_{j,c,t} = s^{2017}_{j,c}$). As shown in Appendix C,

$$s_{j,c,t} - s^{2017}_{j,c} \approx \underbrace{\frac{\bar{s}_{j,c}(1-\bar{s}_{j,c})}{\bar{\alpha}_c}\Delta\alpha_{c,t}}_{\text{intensive margin}} + \underbrace{\frac{\bar{s}_{j,c}(1-\bar{s}_{j,c})}{\bar{\mathbf{1}}_{j,c}}\Delta\mathbf{1}_{j,c,t}}_{\text{extensive margin}} + \underbrace{\text{terms from } c' \neq c}_{\text{cross-country rebalancing}}. \tag{33}$$

In the next section we provide an explicit example of the potential biases that arise

from the endogenous response of these import shares to a tariff.

5 Model Implications

The previous section developed a structural model of endogenous supplier choice in which firms adjust both the intensity of imports from each source country (intensive margin) and the set of countries from which they import (extensive margin) in response to tariff shocks. We now turn to the implications of this framework. Section 5.1 uses the model's predictions to explain the systematic gap between IV and OLS estimates documented in Section 3, showing that endogenous import substitution attenuates the OLS coefficient and that the Bartik IV is immune to this bias. Section 5.2 establishes the link between the empirical exposure measure III_i and the model's primitives and uses this connection to quantify the aggregate PPI effect of the 2018–19 China tariffs. Section 5.3 extends the analysis to a universal tariff counterfactual, motivated by the broad-based tariff policies of 2025–2026, and shows that the aggregate inflationary impact is amplified through broader exposure and higher pass-through.

5.1 Explaining the OLS Bias

Our estimation results show that IV estimates uniformly exceed OLS. For example, in specification (III) for 2017–2019, the (unweighted/weighted) OLS estimates are 0.349/0.313, while the IV estimates are about 0.577/0.531. The IV-to-OLS ratio averages approximately 1.65. To understand the OLS bias, consider a simplified setting with two periods and two countries ($N = 2$). The two periods are a pre-war period ($t = 2017$) and a post-war period ($t = 2018$ and afterwards). We assume that the pre-war period corresponds to the steady state.

The two countries are China and RoW. We assume that China ranks first ($\tilde{b}_1 > \tilde{b}_2$). Moreover, only China ($c = 1$) experiences a tariff increase, while RoW ($c = 2$) faces no change in tariffs. That is,

$$\begin{aligned} d_1^{post} &> d_1^{pre}, \\ d_2^{post} &= d_2^{pre}. \end{aligned}$$

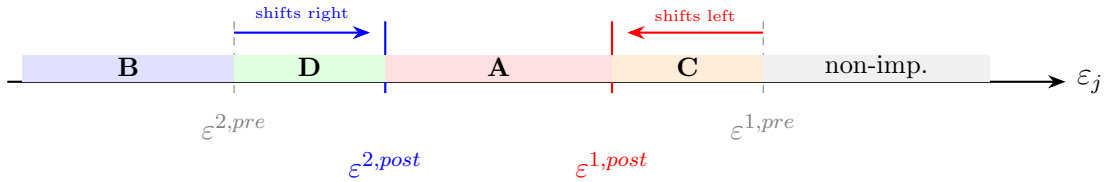
To simplify the analysis (and consistent with our empirical framework), we assume that fixed-cost draws are firm-specific and time-invariant:

$$\epsilon_j^{pre} = \epsilon_j^{post}. \tag{34}$$

That is, we rule out the case where a firm's importing decision changes due to movements in its fixed cost, which isolates the role of the tariff shock.

Consistent with the direct and indirect effects identified in the marginal cost elasticity analysis, the cutoffs move in opposite directions: China's cutoff ε^1 shifts left (decreases), while RoW's cutoff ε^2 shifts right (increases).

Figure 2: Cutoff Movements due to Tariff Hike



NOTE: Dashed lines show pre-war cutoffs; solid lines show post-war cutoffs. China's cutoff shifts left (red), RoW's shifts right (blue), but they do not cross. Type A (stable China-only) survives in the region $(\varepsilon^{2,post}, \varepsilon^{1,post}]$. Type C (exit) drops from the sample.

The movements in the cutoffs partition firms into four types, depending on their pre- and post-war importing status. To streamline the exposition, we focus on the two types that are central to the OLS bias and relegate the full four-type analysis to Appendix D. Type A firms (*stable importers*) have fixed cost draws $\varepsilon^{2,post} < \varepsilon_j \leq \varepsilon^{1,post}$ and import from China in both periods; their import status is unchanged by the tariff. Type C firms (*exit importers*) have draws $\varepsilon^{1,post} < \varepsilon_j \leq \varepsilon^{1,pre}$ and were single-source China importers pre-war but exit importing entirely post-war, because China's cutoff fell below their fixed cost draw. The contrast between these two types captures the extensive margin: type A firms continue importing and experience the tariff as a smooth cost increase, while type C firms lose their import relationship altogether and suffer a discrete cost jump from forfeiting the import variety premium.

We now connect the framework to the standard econometric conditions for OLS bias and IV validity. Define the true structural pass-through model as:

$$\Delta p_j = \beta \, \Delta \tau_j^{IV} + u_j, \tag{35}$$

where $\beta \equiv \text{Cov}(\Delta p_j, \Delta\tau_j^{IV})/\text{Var}(\Delta\tau_j^{IV})$ is the population regression coefficient and $u_j \equiv \Delta p_j - \beta\,\Delta\tau_j^{IV}$ is the structural residual. The parameter β captures the causal effect of tariff exposure (measured at pre-determined weights) on output prices, and u_j captures all other determinants of the price change.

Next, we derive the price change for each firm type. To simplify notation, define $\tilde{\alpha}_1 \equiv \bar{\alpha}_1\gamma_1\phi$, $u_1 \equiv \widehat{e}_{1,t} + \widehat{U}_{1,t} - \widehat{V^*}_{1,t}$ (exogenous aggregate shocks), and $d \equiv \Delta\log(1+d_1)$. A type A firm continues importing from China at the higher tariff rate, so its price change is

$$\Delta p_A = \tilde{\alpha}_1(d + u_1).$$

A type C firm exits importing entirely. Its pre-war marginal cost was $c_j^0/\widetilde{b}_1^{pre}$ (importing from China) and its post-war marginal cost is c_j^0 (domestic sourcing only), so its price change is

$$\Delta p_C = \log c_j^0 - \log(c_j^0/\widetilde{b}_1^{pre}) = \log \widetilde{b}_1^{pre} \equiv \delta_{ext} > 0.$$

The term $\delta_{ext} = \gamma_1\phi\log b_1^{pre}$ is the discrete cost increase from losing the China import variety premium. Importantly, δ_{ext} does not depend on the tariff rate d: once a firm exits importing, its cost is simply c_j^0 regardless of how large the tariff is.[6]

Note that both types had identical pre-war import portfolios (China only), so the Bartik IV assigns both the same exposure: $\Delta\tau^{IV} = d$. However, their OLS regressors diverge: $\Delta\tau^{OLS} = d$ for type A (still importing) but $\Delta\tau^{OLS} = 0$ for type C (no longer in the import data). This divergence is the source of the OLS bias.

The firm's post-war importing decision is an *endogenous variable* that appears inside the OLS regressor through the contemporaneous import shares. Whether a firm continues importing (type A) or exits (type C) depends on whether its fixed cost draw satisfies $\varepsilon_j \leq \varepsilon^{1,post}$. The same draw simultaneously determines the price change—through the variety loss δ_{ext} for exiters—creating a classic simultaneity problem: OLS treats $\Delta\tau_j^{OLS}$ as exogenous, but it is an equilibrium outcome that depends on ε_j.

The direction of the resulting bias is unambiguous. Within the group of type A and type c firms, the firms with the *lowest* OLS regressor (type C, $\Delta\tau^{OLS} = 0$) experience the *largest* price increases (δ_{ext}, the discrete variety loss), while the firms with the

[6]More precisely, $\delta_{ext} = \gamma_1\phi\log b_1^{pre}$ where $b_1^{pre} = [1 + a_1(P_1^{M,pre}/V_1^*)^{-\zeta}]^{1/\zeta}$ reflects the pre-war quality-adjusted cost advantage of Chinese varieties. This equals the full import variety premium that the firm forfeits by exiting.

highest OLS regressor (type A, $\Delta\tau^{OLS} = d$) experience *moderate* price increases ($\tilde{\alpha}_1 d$, the smooth tariff pass-through). This negative correlation between the OLS regressor and the structural residual biases the OLS coefficient downward.

5.2 Aggregate PPI Inflation: China Tariffs

We now connect the structural model to the aggregate inflationary impact of the 2018–19 China tariffs. We first establish the link between the empirical exposure measure III_i and the model's primitives, and then use the estimated pass-through coefficient to quantify the effect on aggregate producer price inflation.

Recall that the empirical imported input intensity, III_i, measures the fraction of total variable cost attributable to imported intermediates. Our model provides a structural counterpart. Summing over all source countries and dividing by total variable cost:[7]

$$III_j^{\text{model}} = \frac{\sum_c P_{c,t}^M M_{j,c,t}}{TVC_{j,t}} = \phi \sum_{c=1}^{N} \gamma_c \, \alpha_{c,t} \, \mathbf{1}_{j,c,t} \equiv \phi \, m_{j,t}, \tag{36}$$

where $m_{j,t}$ denotes the import intensity within intermediates.

Recall the structural import share from equation (19):

$$s_{j,c,t} = \frac{\gamma_c \, \alpha_{c,t} \, \mathbf{1}_{j,c,t}}{\sum_{c'=1}^{N} \gamma_{c'} \, \alpha_{c',t} \, \mathbf{1}_{j,c',t}} = \frac{\gamma_c \, \alpha_{c,t} \, \mathbf{1}_{j,c,t}}{m_{j,t}}. \tag{37}$$

Combining (36) and (37) yields the key identity:

$$III_j^{model} \cdot s_{j,c,t} = \phi \, \gamma_c \, \alpha_{c,t} \, \mathbf{1}_{j,c,t}. \tag{38}$$

The product $III_j^{model} \times s_{j,c,t}$ equals the fraction of the firm's *total variable cost* attributable to imports from country c. This decomposes cleanly into three structural primitives: the intermediate cost share ϕ, the Cobb–Douglas weight γ_c, and the CES import expenditure share $\alpha_{c,t}$ (multiplied by the importing indicator). The identity

[7] The Cobb–Douglas production function (9) implies that total intermediate spending is a fixed fraction ϕ of total variable cost, $\tilde{V}_t \, X_{j,t} = \phi \cdot TVC_{j,t}$. Within intermediates, the Cobb–Douglas aggregation across types implies that spending on type c is a fraction γ_c of total intermediate spending. Within type c, the CES composite between domestic and foreign varieties implies that import spending is a fraction $\alpha_{c,t}$ of type-c spending, conditional on importing ($\mathbf{1}_{j,c,t} = 1$). Combining these cost shares: $P_{c,t}^M \, M_{j,c,t} = \gamma_c \, \alpha_{c,t} \, \mathbf{1}_{j,c,t} \cdot \tilde{V}_t \, X_{j,t} = \phi \, \gamma_c \, \alpha_{c,t} \, \mathbf{1}_{j,c,t} \cdot TVC_{j,t}$.

has a practical implication: it shows that the regressor in the interacted IV specification, $III_i \times \Delta\tau_j^{IV} = \sum_c III_j^{model}\, s_{j,c}^{2017}\, \Delta\log(1+d_c)$, is the empirical counterpart of the model's cost-base-weighted tariff exposure.

Quantifying the Aggregate PPI Effect

Let ω_j denote firm j's weight in the aggregate PPI (reflecting its share of total shipments). Aggregate PPI inflation from the China tariff is:

$$\Delta\log P^{PPI} = \sum_j \omega_j\, \Delta\log p_j = \hat{\beta}^{IV} \times \Delta\log(1+d_{\text{China}}) \times \underbrace{\sum_j \omega_j\, III_j\, s_{j,\text{China}}^{2017}}_{\equiv\, \overline{III\cdot s}}. \quad (39)$$

The aggregate exposure statistic $\overline{III \cdot s}$ is the PPI-weighted average of $III_j \times s_{j,\text{China}}^{2017}$ across *all* firms in the PPI—including non-importers, for whom $III_j = s_{j,c} = 0$. By identity (38), this equals the PPI-weighted average of $\phi\, \gamma_{\text{China}}\, \alpha_{\text{China}}\, \mathbf{1}_{j,\text{China}}$, i.e., the average share of total variable costs from Chinese imports across the entire manufacturing sector.

As a preliminary estimate, we use $\overline{III}^{\text{imp}} \approx 0.20$ and $\bar{s}^{\text{imp}} \approx 0.25$ among importers and suppose that importers account for roughly $\theta_{\text{imp}} \approx 0.60$ of PPI-weighted manufacturing output.Noting that non-importers contribute zero, we obtain $\overline{III \cdot s} \approx 0.60 \times 0.20 \times 0.25 = 0.03$. For a 15 percentage point increase in the China tariff ($\Delta d_{\text{China}} = 0.15$):

$$\Delta\log P^{PPI,A} \approx 0.577 \times 0.140 \times 0.03 \approx 0.24 \text{ pp}. \quad (40)$$

This implies that the 2018–19 China tariffs raised aggregate manufacturing PPI by approximately 0.2–0.3 percentage points through the input cost channel alone, abstracting from the competition (output tariff) channel estimated separately in Section 3.

5.3 Counterfactual: Universal Tariff

We now consider a counterfactual in which all N supplier countries simultaneously face the same ad valorem tariff increase. This exercise is directly motivated by recent

U.S. trade policy developments.[8] These policy developments mark a fundamental shift from the *targeted* tariffs of the 2018–19 trade war—which were concentrated on Chinese imports and specific product categories—to *broad-based* tariffs applied across virtually all supplier countries. Our model provides a natural framework for analyzing how this shift amplifies the inflationary impact of tariffs on domestic producer prices. The key insight is that the aggregate PPI effect of a universal tariff is *more than proportional* to the share of imports tariffed, because a universal tariff simultaneously shuts down the substitution channels that cushion the impact of a targeted tariff.

To fix ideas, we conduct the counterfactual at a common rate of $\Delta d = 15\%$, which is within the range of the tariff rates imposed or proposed during 2025–2026. Under a universal tariff $d_{c,t} = d_t$ for all c, the chain rule therefore yields:

$$\frac{d \log \mathbb{E}[MC_{j,t}]}{d \log(1 + d_t)} = \sum_{c=1}^{N} \frac{d \log \mathbb{E}[MC_{j,t}]}{d \log(1 + d_{c,t})}. \tag{41}$$

The universal tariff elasticity of expected marginal cost is the sum of the country-specific elasticities. Since each country-specific elasticity is strictly positive—the direct effect dominates the crowding-out effect for every c, as shown by the Corollary in Section 4—the universal elasticity is unambiguously positive and strictly exceeds any individual country-specific elasticity.

Aggregate PPI Under Universal Tariff

Under a universal tariff of $\Delta d = 15\%$, every importing firm faces full tariff exposure. The IV exposure becomes:

$$\Delta \tau_j^{IV,U} = \sum_c s_{j,c}^{2017} \, \Delta \log(1.15) = \log(1.15), \tag{42}$$

[8] On April 2, 2025, President Trump invoked the International Emergency Economic Powers Act (IEEPA) to impose a universal baseline tariff of 10% on imports from nearly all countries, with higher "reciprocal" rates of 11–50% on 57 trading partners deemed to have nonreciprocal practices. Although the Supreme Court struck down the IEEPA-based tariffs in *Learning Resources, Inc. v. Trump* on February 20, 2026, the administration immediately reimposed a universal 10% tariff under Section 122 of the Trade Act of 1974 for 150 days. The administration has also signaled interest in establishing a permanent universal tariff baseline.

where the last equality uses $\sum_c s_{j,c}^{2017} = 1$. The aggregate PPI effect at first order is:

$$\Delta \log P^U = \hat{\beta}^{IV} \times \log(1.15) \times \overline{III}, \tag{43}$$

where $\overline{III} = \sum_j \omega_j \, III_j$ is now understood as the economy-wide PPI-weighted average including non-importers: $\overline{III} = \theta_{\text{imp}} \times \overline{III}^{\text{imp}} \approx 0.60 \times 0.20 = 0.12$. Comparing with (39):

$$\frac{\Delta \log P^U}{\Delta \log P} = \frac{\overline{III}}{\overline{III \cdot s}} = \frac{\overline{III}^{\text{imp}}}{\overline{III \cdot s}^{\text{imp}}} \approx 4, \tag{44}$$

where the second equality uses the fact that θ_{imp} cancels in the ratio.

At first order, the universal 15% tariff raises aggregate PPI by approximately $0.577 \times 0.140 \times 0.12 \approx 1.0$ percentage point—roughly four times the China-only effect of ≈ 0.24 pp. However, the first-order calculation treats the pass-through coefficient $\hat{\beta}^{IV}$ as invariant across scenarios. The structural model implies that this understates the true universal tariff effect, for the following reason.

6 Conclusion

This paper has provided new firm-level evidence on the pass-through from import tariffs to U.S. producer prices. Using confidential LFTTD data linked with BLS producer price microdata, we construct firm-specific measures of effective input tariff exposure and estimate the causal effect of tariff changes on output prices using Bartik-style instrumental variables. Our IV estimates imply a pass-through of approximately 0.5–0.6 at a 12-month horizon, meaning that roughly half of a tariff increase on imported inputs is reflected in higher domestic output prices within a year.

Several findings deserve emphasis. First, the pass-through is economically meaningful but well below unity, indicating that firms absorb a substantial portion of tariff costs into their margins—at least when measured at the industry level of exposure. Second, IV estimates are systematically larger than OLS, consistent with our formal demonstration that endogenous import substitution attenuates the OLS coefficient toward zero. Our structural model of endogenous supplier choice provides a micro-foundation for this gap. The model shows that firms respond to tariff increases through two channels: an intensive margin (CES substitution from imported toward domestic intermediates within existing supplier relationships) and an extensive

margin (dropping high-tariff suppliers and switching to alternatives). Both channels compress the contemporaneous effective tariff relative to the pre-determined Bartik IV, attenuating the OLS coefficient.

Third, using the estimated pass-through and the distribution of firm-level tariff exposure, we find that the 2018–19 China tariffs raised aggregate manufacturing PPI by approximately 0.2–0.3 percentage points through the input cost channel. This estimate captures only the direct cost channel and does not include the competition (output tariff) channel, general equilibrium effects on wages and domestic input prices, or second-round effects through which higher PPI feeds into consumer prices.

Fourth, we use the model to analyze a counterfactual that is directly motivated by the shift in U.S. trade policy during 2025–2026: a universal tariff applied simultaneously to all supplier countries. Our model predicts that a universal tariff of the same rate generates an aggregate PPI effect that is 4–5 times larger than a China-only tariff. The per-unit pass-through is higher, because the substitution "escape valve"—the ability to shift sourcing toward untariffed countries—is simultaneously shut down across all supplier countries.

These findings have clear policy implications. The welfare costs of import tariffs extend well beyond the direct burden on importers documented by Amiti, Redding, and Weinstein (2019): tariffs on intermediate inputs propagate through the production chain to raise domestic output prices, amplifying the aggregate price effects of trade protection. For monetary policymakers calibrating their response to tariff shocks, our results suggest that the full inflationary impact occurs within one year, implying a relatively rapid response of prices to trade policy announcements.

Our analysis opens several avenues for future work. First, the heterogeneity of pass-through across firms of different sizes, market power, and input substitution elasticities deserves further investigation. Second, tracing the chain one step further—from producer prices to consumer prices—would complete the welfare accounting. Third, the interaction between tariff pass-through and other cost shocks (e.g., supply chain disruptions during COVID-19) could illuminate how firms manage simultaneous input cost pressures. Fourth, extending the structural model to incorporate general equilibrium effects—including wage adjustments, domestic input price changes, and demand-side responses—would provide a more complete picture of the aggregate inflationary impact of tariffs, particularly in the context of the broad-based tariff policies that have become a central feature of contemporary U.S. trade policy.

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Supplementary Material

— For Online Publications Only —

A Merging PPI and LFTTD Databases

We merge the PPI and LFTTD databases using a multi-step name-matching algorithm designed to maximize match rates while minimizing false positives. Because the two databases do not share a common firm identifier, we rely on business names as the primary matching variable.[9]

The matching procedure proceeds in several stages. First, we conduct an iterative word match routine that is repeated on two versions of the business name: (i) the raw original text from the PPI records and (ii) a cleaned version of the name (converted to uppercase, with punctuation such as "&" and "-" removed, and common business suffixes such as "INC," "CO," "LLC," "CORP," and "LTD" standardized and dropped).

The word match routine is carried out in three steps. In the first step, we directly match the full cleaned name string against the LFTTD database. This step captures exact or near-exact name matches. In the second step, the name string is tokenized into its constituent words, and we attempt to match the first five words of the name against the LFTTD. This step captures cases where business names differ only in trailing words (e.g., division names or location identifiers). The third step is identical to the second step except that we use a maximum of four words only, capturing additional matches where names share a common root but differ in specificity.

After exhausting the deterministic matching steps, we apply a fuzzy name matching procedure to the remaining unmatched records. Specifically, we use the SAS DQMATCH function with a sensitivity parameter of 95 to generate a standardized version of the cleaned business name. We retain matches to a unique firm identifier. For cases in which a PPI business name matches to multiple LFTTD firm identifiers, we resolve the ambiguity by matching the name strings word by word and retaining only those matches for which at least two constituent words agree, starting from the maximum possible number of matched words.

[9] In constructing our sample from the LFTTD, we require: (i) positive and nonmissing import or export value; (ii) nonmissing shipping costs and duties; (iii) nonmissing firm identifier; and (iv) we drop warehouse entries.

To further validate the matches, we impose several consistency checks. We require that matched firms belong to the same broad industry group (at the 2-digit NAICS level) and, where geographic information is available, that they are located in the same state. These filters help eliminate spurious matches arising from common business names (e.g., "American Manufacturing" or "National Products").

The resulting matched sample contains approximately 6,000 unique manufacturing firms that appear in both the PPI and LFTTD databases during the 2005–2022 period. These firms account for a substantial share of total U.S. manufacturing imports, suggesting that our sample is representative of the population of importing manufacturers.

B Empirical Results using Alternative IV

As a robustness check, we also construct an instrument using de facto (effectively collected) tariff rates:

$$\Delta \log \left(1 + d_{j,t}^{\text{IV,eff}}\right) \equiv \sum_{(h,c)} \frac{v_{(h,c)\to j}^{(2017)}}{\sum_{(h,c)} v_{(h,c)\to j}^{(2017)}} \Delta \log \left(1 + d_{h,c,t}^{eff}\right). \tag{B-1}$$

Table B-1 reports 2SLS estimates using the de facto Bartik IV based on Specifications (I) and (II). The results closely parallel the de jure IV: 0.211 (0.035) versus 0.203 (0.035) in specification (I) unweighted, and 0.152 (0.034) versus 0.144 (0.033) in specification (II). The first-stage is even stronger, with coefficients of 0.82 and F-statistics exceeding 830. The near-identical second-stage estimates across the two IV constructions—despite the de jure and de facto tariff rates differing by the extent of product exclusions and administrative exemptions—provide reassuring evidence that our findings are not sensitive to the specific treatment of tariff exceptions.

Table B-2 confirms these patterns using the de facto Bartik IV. In specification (III), the IV estimate is 0.600 (0.145) unweighted and 0.544 (0.143) weighted—very close to the de jure results of 0.577 and 0.531. In specification (IV), the estimates are 0.589 (0.151) and 0.497 (0.160). The first-stage is stronger than for the de jure instrument, with coefficients of 0.76 and F-statistics exceeding 450. The robust agreement between the de jure and de facto IV results across all specifications reinforces the conclusion that our pass-through estimates are not artifacts of how tariff exemptions are measured.

Table B-1: IV Regression Results, De Facto Tariffs

Panel A: Second-stage results	Specification (I)		Specification (II)	
	Unwt.	Wt.	Unwt.	Wt.
$\Delta \log(1 + d^{\text{input,eff}})$	0.211	0.201	0.152	0.127
	(0.035)	(0.036)	(0.034)	(0.037)
$\Delta \log(1 + d^{\text{output,eff}})$	0.242	0.267		
	(0.070)	(0.074)		
Observations	296k	296k	316k	316k
Firms	3,000	3,000	3,200	3,200
Panel B: First-stage results				
$\Delta \log(1 + d^{\text{IV,eff}})$	0.819	0.819	0.815	0.816
	(0.028)	(0.027)	(0.025)	(0.024)
R^2	0.700	0.695	0.748	0.748
F-statistic	830.6	931.6	1039.0	1160.0
Firm FE	Yes	Yes	Yes	Yes
Industry×Month FE	No	No	Yes	Yes

Notes: Panel A reports 2SLS estimates using the de facto Bartik IV from equation (B-1). Panel B reports first-stage results. Standard errors clustered at the firm level. F-statistic is the Kleibergen-Paap weak-instrument test.

The first-stage regressions in Tables 5 and B-2 display two systematic patterns. First, the first-stage coefficients are uniformly below one (0.53–0.76), confirming the structural prediction that endogenous import substitution compresses the contemporaneous effective tariff relative to the pre-determined Bartik instrument. Second, the de facto IV yields larger first-stage coefficients (0.75–0.76) than the de jure IV (0.65–0.67), consistent with the de facto measure absorbing less measurement noise because it reflects actually collected duties rather than announced statutory rates.

C Proofs

Proof of Proposition 1. First, we consider the firm's importing decision and derive the within-type Unit Cost $\tilde{V}_{c,t}$. Consider the type-c composite intermediate, which aggregates domestic and foreign varieties via the CES technology (11). The firm minimizes the cost of producing one unit of $X_{j,c,t}$, taking as given the price of

Table B-2: IV with Exposure Interaction, De Facto Tariffs

Panel A: Second-stage results	Specification (III)		Specification (IV)	
	Unwt.	Wt.	Unwt.	Wt.
$III \times \Delta \log(1 + d^{\text{input,eff}})$	0.600	0.544	0.589	0.497
	(0.145)	(0.143)	(0.151)	(0.160)
$IS \times \Delta \log(1 + d^{\text{output,eff}})$	0.604	0.690		
	(0.187)	(0.199)		
Obs.	296k	296k	296k	296k
Firms	3,000	3,000	3,000	3,000
Panel B: First-stage results				
$III \times \Delta \log(1 + d^{\text{IV,eff}})$	0.761	0.763	0.748	0.749
	(0.035)	(0.034)	(0.035)	(0.032)
R^2	0.712	0.708	0.753	0.753
F-statistic	461.6	507.1	458.1	532.2
Firm FE	Yes	Yes	Yes	Yes
Ind.×Month FE	No	No	Yes	Yes

Notes: Panel A reports 2SLS estimates using the de facto Bartik IV from equation (B-1). Panel B reports first-stage results. Standard errors clustered at the firm level. F-statistic is the Kleibergen-Paap weak-instrument test.

the domestic variety $V^*_{c,t}$ and the landed import price $P^M_{c,t}$. The CES has elasticity of substitution $\sigma = 1 + \zeta$ between domestic and foreign varieties.

Normalize $X_{j,c,t} = 1$ to derive the unit cost:

$$\min_{Z,M} \; V^*_{c,t}\, Z + P^M_{c,t}\, M \quad \text{s.t.} \quad \left[Z^{\zeta/(1+\zeta)} + a_c^{1/(1+\zeta)}\, M^{\zeta/(1+\zeta)}\right]^{(1+\zeta)/\zeta} = 1.$$

Setting up the Lagrangian $\mathcal{L} = V^*_{c,t}\, Z + P^M_{c,t}\, M - \lambda \left[\left(Z^{\zeta/(1+\zeta)} + a_c^{1/(1+\zeta)}\, M^{\zeta/(1+\zeta)}\right)^{(1+\zeta)/\zeta} - 1\right]$, the first-order conditions are:

$$V^*_{c,t} = \lambda \cdot Z^{-1/(1+\zeta)} \cdot X^{1/\zeta}_{j,c,t}, \tag{C-1}$$

$$P^M_{c,t} = \lambda \cdot a_c^{1/(1+\zeta)}\, M^{-1/(1+\zeta)} \cdot X^{1/\zeta}_{j,c,t}. \tag{C-2}$$

From (C-1): $Z = (\lambda / V^*_{c,t})^{1+\zeta} \cdot X^{(1+\zeta)/\zeta}_{j,c,t}$. From (C-2): $M = a_c\, (\lambda / P^M_{c,t})^{1+\zeta} \cdot X^{(1+\zeta)/\zeta}_{j,c,t}$. Substituting both into the CES constraint ($X_{j,c,t} = 1$) yields $\lambda = \left[(V^*_{c,t})^{-\zeta} + a_c\, (P^M_{c,t})^{-\zeta}\right]^{-1/\zeta}$,

which is the shadow price of the constraint $X_{j,c,t} = 1$, and equals the unit cost:

$$\tilde{V}_{c,t} = \lambda = \left[(V^*_{c,t})^{-\zeta} + a_c \left(P^M_{c,t}\right)^{-\zeta}\right]^{-1/\zeta}. \tag{C-3}$$

This is the unit cost of the type-c composite when the firm imports from country c. If the firm does *not* import from c (i.e., $c \notin \mathcal{J}_{j,t}$), then $M_{j,c,t} = 0$ and the unit cost is simply $V^*_{c,t}$.

If the firm imports from c (i.e., $c \in \mathcal{J}_{j,t}$), then

$$\tilde{V}_{c,t} = \left[(V^*_{c,t})^{-\zeta}\left(1 + a_c\left(\frac{P^M_{c,t}}{V^*_{c,t}}\right)^{-\zeta}\right)\right]^{-1/\zeta} = V^*_{c,t} \cdot \left[1 + a_c\left(\frac{P^M_{c,t}}{V^*_{c,t}}\right)^{-\zeta}\right]^{-1/\zeta} = \frac{V^*_{c,t}}{b_{c,t}},$$

where the last equality uses $b_{c,t} = [1 + a_c(P^M_{c,t}/V^*_{c,t})^{-\zeta}]^{1/\zeta}$ from (17). Therefore importing from country c reduces the unit cost of type-c intermediates from $V^*_{c,t}$ to $V^*_{c,t}/b_{c,t}$, confirming the interpretation of $b_{c,t} > 1$ as a cost-reduction factor. Summarizing:

$$\tilde{V}_{c,t} = \begin{cases} V^*_{c,t}/b_{c,t} & \text{if } c \in \mathcal{J}_{j,t} \text{ (firm imports from } c\text{)}, \\ V^*_{c,t} & \text{if } c \notin \mathcal{J}_{j,t} \text{ (firm does not import from } c\text{)}. \end{cases} \tag{C-4}$$

Second, we derive the aggregate unit cost $\tilde{V}_t$. The aggregate intermediate bundle combines N types via the Cobb-Douglas aggregator (10): $X_{j,t} = \prod_{c=1}^{N} X_{j,c,t}^{\gamma_c}$. The firm chooses $\{X_{j,c,t}\}_{c=1}^{N}$ to minimize total intermediate cost $\sum_c \tilde{V}_{c,t} X_{j,c,t}$ subject to $\prod_c X_{j,c,t}^{\gamma_c} = X_{j,t}$. The Lagrangian first-order condition for type c is:

$$\tilde{V}_{c,t} = \mu \cdot \gamma_c \frac{X_{j,t}}{X_{j,c,t}},$$

where μ is the Lagrange multiplier. This immediately yields the Cobb-Douglas expenditure share property:

$$\tilde{V}_{c,t} X_{j,c,t} = \gamma_c \mu X_{j,t} = \gamma_c \cdot \tilde{V}_t X_{j,t}, \tag{C-5}$$

where $\tilde{V}_t \equiv \mu$ is the unit cost of the aggregate bundle. Thus the expenditure share on each type equals its Cobb-Douglas weight γ_c, regardless of prices. From the first-order condition: $X_{j,c,t} = \gamma_c \tilde{V}_t X_{j,t}/\tilde{V}_{c,t}$. Substituting into the Cobb-Douglas constraint

yields:

$$\tilde{V}_t = \prod_{c=1}^{N} \left(\frac{\tilde{V}_{c,t}}{\gamma_c} \right)^{\gamma_c}. \tag{C-6}$$

Splitting the product over the importing set $\mathcal{J}_{j,t}$ and its complement, and using (C-4):

$$\tilde{V}_t = \prod_{c \in \mathcal{J}_{j,t}} \left(\frac{V_{c,t}^* / b_{c,t}}{\gamma_c} \right)^{\gamma_c} \cdot \prod_{c \notin \mathcal{J}_{j,t}} \left(\frac{V_{c,t}^*}{\gamma_c} \right)^{\gamma_c} = \underbrace{\prod_{c=1}^{N} \left(\frac{V_{c,t}^*}{\gamma_c} \right)^{\gamma_c}}_{\equiv V_t^*} \cdot \underbrace{\prod_{c \in \mathcal{J}_{j,t}} b_{c,t}^{-\gamma_c}}_{\equiv 1/B_{j,t}},$$

where:

$$V_t^* \equiv \prod_{c=1}^{N} \left(\frac{V_{c,t}^*}{\gamma_c} \right)^{\gamma_c}, \qquad B_{j,t} \equiv \prod_{c \in \mathcal{J}_{j,t}} b_{c,t}^{\gamma_c}. \tag{C-7}$$

Here V_t^* is the unit cost for a non-importing firm and $B_{j,t} \geq 1$ is the cost-reduction factor from importing. Therefore:

$$\tilde{V}_t = \frac{V_t^*}{B_{j,t}}. \tag{C-8}$$

Lastly, we derive the firm's production decision and derive its overall unit cost from production. Starting from the production function $Y_{j,t} = \Omega_{j,t} X_{j,t}^{\phi} L_{j,t}^{1-\phi}$, the firm minimizes:

$$\min_{X_{j,t}, L_{j,t}} \tilde{V}_t X_{j,t} + W_t L_{j,t} \quad \text{s.t.} \quad Y_{j,t} = \Omega_{j,t} X_{j,t}^{\phi} L_{j,t}^{1-\phi}, \tag{C-9}$$

where $\tilde{V}_t$ is the effective unit cost of the composite intermediate input (which depends on the importing set $\mathcal{J}_{j,t}$). Setting up the Lagrangian:

$$\mathcal{L} = \tilde{V}_t X_{j,t} + W_t L_{j,t} + \lambda \left[Y_{j,t} - \Omega_{j,t} X_{j,t}^{\phi} L_{j,t}^{1-\phi} \right].$$

The first-order conditions are:

$$\frac{\partial \mathcal{L}}{\partial X_{j,t}} = \tilde{V}_t - \lambda \phi \Omega_{j,t} X_{j,t}^{\phi-1} L_{j,t}^{1-\phi} = 0 \implies \tilde{V}_t = \lambda \phi \frac{Y_{j,t}}{X_{j,t}}, \tag{C-10}$$

$$\frac{\partial \mathcal{L}}{\partial L_{j,t}} = W_t - \lambda (1-\phi) \Omega_{j,t} X_{j,t}^{\phi} L_{j,t}^{-\phi} = 0 \implies W_t = \lambda (1-\phi) \frac{Y_{j,t}}{L_{j,t}}. \tag{C-11}$$

Dividing (C-10) by (C-11) yields the optimal input ratio:

$$\frac{X_{j,t}}{L_{j,t}} = \frac{\phi}{1-\phi} \cdot \frac{W_t}{\tilde{V}_t}. \tag{C-12}$$

From (C-10): $X_{j,t} = \lambda\phi Y_{j,t}/\tilde{V}_t$. From (C-11): $L_{j,t} = \lambda(1-\phi)Y_{j,t}/W_t$. Substituting into the production function yields $\lambda = \frac{1}{\Omega_{j,t}} \left(\frac{\tilde{V}_t}{\phi}\right)^{\phi} \left(\frac{W_t}{1-\phi}\right)^{1-\phi}$, which is the marginal cost. Total variable cost is:

$$TVC = \tilde{V}_t X_{j,t} + W_t L_{j,t} = \lambda Y_{j,t} = \frac{Y_{j,t}}{\Omega_{j,t}} \left(\frac{\tilde{V}_t}{\phi}\right)^{\phi} \left(\frac{W_t}{1-\phi}\right)^{1-\phi}.$$

□

Derivation of Equations (18) and (19). From the Cobb-Douglas aggregation of intermediate types (10), the standard property of the Cobb-Douglas function implies that the cost share of type-c intermediates in total intermediate spending $\tilde{V}_t X_{j,t}$ is exactly γ_c. That is, if $\tilde{V}_{c,t}$ denotes the unit cost of the type-c composite and $\tilde{V}_t$ the unit cost of the aggregate intermediate bundle, then:

$$\tilde{V}_{c,t}\, X_{j,c,t} = \gamma_c\, \tilde{V}_t\, X_{j,t}.$$

Within each type c, the CES composite (11) combines domestic variety $Z_{j,c,t}$ and foreign variety $M_{j,c,t}$. If firm j imports from country c (i.e., $c \in \mathcal{J}_{j,t}$), the expenditure share of the foreign variety in type-c spending is $\alpha_{c,t}$ from (31):

$$P^M_{c,t}\, M_{j,c,t} = \alpha_{c,t} \cdot \tilde{V}_{c,t}\, X_{j,c,t}.$$

If firm j does not import from country c ($c \notin \mathcal{J}_{j,t}$), then $M_{j,c,t} = 0$ and import spending from c is zero. Combining both cases using the indicator $\mathbf{1}_{j,c,t} \equiv \mathbf{1}[c \in \mathcal{J}_{0,j,t}]$:

$$P^M_{c,t}\, M_{j,c,t} = \alpha_{c,t}\, \mathbf{1}_{j,c,t} \cdot \tilde{V}_{c,t}\, X_{j,c,t}.$$

Substituting the Cobb-Douglas expenditure identity from Step 1:

$$P^M_{c,t}\, M_{j,c,t} = \gamma_c\, \alpha_{c,t}\, \mathbf{1}_{j,c,t} \cdot \tilde{V}_t\, X_{j,t}.$$

The value of firm j's imports from country c equals the Cobb-Douglas weight γ_c, times the CES import expenditure share $\alpha_{c,t}$, times the importing indicator $\mathbf{1}_{j,c,t}$, times total intermediate spending $\tilde{V}_t X_{j,t}$. Total import spending is thus given by $\sum_{c'=1}^{N} P_{c',t}^{M} M_{j,c',t} = \left(\sum_{c'=1}^{N} \gamma_{c'} \alpha_{c',t} \mathbf{1}_{jc',t}\right) \cdot \tilde{V}_t X_{j,t}$. The previous two equations imply (19). □

Proof of Proposition 3. Taking expectations over ϵ, we can write the expected variable cost below:

$$\begin{aligned}
\mathbb{E}[TVC_{j,t}] &= \left(1 - F\left(\epsilon_{j,t}^{1}\right)\right) TVC_{j,t}^{0} + \sum\nolimits_{k=1}^{N} \left(F\left(\epsilon_{j,t}^{k}\right) - F\left(\epsilon_{j,t}^{k+1}\right)\right) TVC_{j,t}^{k} \\
&= TVC_{j,t}^{0} + \sum\nolimits_{k=1}^{N} F\left(\epsilon_{j,t}^{k}\right) \left(TVC_{j,t}^{k} - TVC_{j,t}^{k-1}\right) \\
&= TVC_{j,t}^{0} - \sum\nolimits_{k=1}^{N} F\left(\epsilon_{j,t}^{k}\right) \left(\tilde{b}_{k,t} - 1\right) TVC_{j,t}^{k} \\
&= TVC_{j,t}^{0} - \sum\nolimits_{k=1}^{N} F\left(\epsilon_{j,t}^{k}\right) \epsilon_{j,t}^{k} C_{F,t}.
\end{aligned}$$

On the other hand, consider the expected value of the fixed costs as

$$\begin{aligned}
\mathbb{E}[TFC_{j,t}] &= \sum\nolimits_{k=1}^{N} \left(F\left(\epsilon_{j,t}^{k}\right) - F\left(\epsilon_{j,t}^{k+1}\right)\right) \mathbb{E}\left(\epsilon | \, \epsilon_{j,t}^{k+1} \leq \epsilon < \epsilon_{j,t}^{k}\right) \left(kC_{F,t}\right) \\
&= \left(\int_{\epsilon_{j,t}^{2}}^{\epsilon_{j,t}^{1}} \epsilon dF\left(\epsilon\right) + 2\int_{\epsilon_{j,t}^{3}}^{\epsilon_{j,t}^{2}} \epsilon dF\left(\epsilon\right) + \cdots + N\int_{0}^{\epsilon_{j,t}^{N}} \epsilon dF\left(\epsilon\right)\right) C_{F,t} \\
&= \sum\nolimits_{k=1}^{N} k \left(\int_{0}^{\epsilon_{j,t}^{k}} \epsilon dF\left(\epsilon\right) - \int_{0}^{\epsilon_{j,t}^{k+1}} \epsilon dF\left(\epsilon\right)\right) C_{F,t} \\
&= \left(\sum\nolimits_{k=1}^{N} k \int_{0}^{\epsilon_{j,t}^{k}} \epsilon dF\left(\epsilon\right) - \sum\nolimits_{k=1}^{N-1} k \int_{0}^{\epsilon_{j,t}^{k+1}} \epsilon dF\left(\epsilon\right)\right) C_{F,t} \\
&= \left(\sum\nolimits_{k=1}^{N} k \int_{0}^{\epsilon_{j,t}^{k}} \epsilon dF\left(\epsilon\right) - \sum\nolimits_{k=2}^{N} \left(k-1\right) \int_{0}^{\epsilon_{j,t}^{k}} \epsilon dF\left(\epsilon\right)\right) C_{F,t} \\
&= C_{F,t} \sum\nolimits_{k=1}^{N} \int_{0}^{\epsilon_{j,t}^{k}} \epsilon dF\left(\epsilon\right).
\end{aligned}$$

So taking expectations over the fixed cost draw, the expected total cost is:

$$\begin{aligned}
\mathbb{E}[TC_{j,t}] &= \mathbb{E}[TVC_{j,t}] + \mathbb{E}[TFC_{j,t}] \qquad \text{(C-13)}\\
&= TVC^0_{j,t} - \sum\nolimits_{k=1}^{N} F\left(\epsilon^k_{j,t}\right)\epsilon^k_{j,t} C_{F,t} + C_{F,t}\sum\nolimits_{k=1}^{N}\int_0^{\epsilon^k_{j,t}} \epsilon dF\left(\epsilon\right)\\
&= TVC^0_{j,t} - C_{F,t}\sum\nolimits_{k=1}^{N}\int_0^{\epsilon^k_{j,t}} \left(\epsilon^k_{j,t} - \epsilon\right) dF\left(\epsilon\right)\\
&= TVC^0_{j,t} - C_{F,t}\sum\nolimits_{k=1}^{N} G(\varepsilon^k_{j,t}),
\end{aligned}$$

where $TVC^0_{j,t} \equiv \frac{C^*_t}{\Omega_{j,t}} Y_{j,t}$ denotes the total variable cost for a non-importing firm, and

$$G(\varepsilon^k_{j,t}) = \int_0^{\varepsilon^k_{j,t}} (\varepsilon^k_{j,t} - \varepsilon)\, dF(\varepsilon) \qquad \text{(C-14)}$$

captures the expected cost savings from the option to import variety k.

The expected marginal cost with respect to output is given by

$$\begin{aligned}
\mathbb{E}[MC_{j,t}] &\equiv \frac{d\mathbb{E}[TC_{j,t}]}{dY_{j,t}} = \frac{dTVC^0_{j,t}}{dY_{j,t}} - C_{F,t}\sum\nolimits_{k=1}^{N}\frac{dG(\varepsilon^k_{j,t})}{dY_{j,t}}\\
&= \frac{TVC^0_{j,t}}{Y_{j,t}} - C_{F,t}\sum\nolimits_{k=1}^{N} F(\varepsilon^k_{j,t})\frac{d\varepsilon^k_{j,t}}{dY_{j,t}}\\
&= \frac{TVC^0_{j,t}}{Y_{j,t}} - C_{F,t}\sum\nolimits_{k=1}^{N} F(\varepsilon^k_{j,t})\frac{\varepsilon^k_{j,t}}{Y_{j,t}}\\
&\equiv c^0_{j,t} - c^f_{j,t}\sum\nolimits_{k=1}^{N}\varepsilon^k_{j,t} F(\varepsilon^k_{j,t}),
\end{aligned}$$

where we have used the results $\frac{dTVC^0_{j,t}}{dY_{j,t}} = \frac{TVC^0_{j,t}}{Y_{j,t}}$, $\frac{d\varepsilon^k_{j,t}}{dY_{j,t}} = \frac{\varepsilon^k_{j,t}}{Y_{j,t}}$, and $G'\left(\varepsilon^k_{j,t}\right) = F\left(\varepsilon^k_{j,t}\right)$. □

Derivation of Equation (28). Define the relative import price $r_{c,t} \equiv P^M_{c,t}/V^*_{c,t}$. The derivation proceeds via the chain rule:

$$\frac{d\tilde{b}_{c,t}}{dd_{c,t}} = \frac{d\tilde{b}_{c,t}}{db_{c,t}} \cdot \frac{db_{c,t}}{dr_{c,t}} \cdot \frac{dr_{c,t}}{dP^M_{c,t}} \cdot \frac{dP^M_{c,t}}{dd_{c,t}}.$$

First, from $P^M_{c,t} = e_t\, U_{c,t}\, \tau_{c,t}\, (1 + d_{c,t})$:

$$\frac{dP^M_{c,t}}{dd_{c,t}} = e_t\, U_{c,t}\, \tau_{c,t} = \frac{P^M_{c,t}}{1 + d_{c,t}}. \tag{C-15}$$

Second, since $V^*_{c,t}$ is independent of $d_{c,t}$:

$$\frac{dr_{c,t}}{dP^M_{c,t}} = \frac{1}{V^*_{c,t}}. \tag{C-16}$$

Third, from $b_{c,t} = [1 + a_c\, r^{-\zeta}_{c,t}]^{1/\zeta}$, we have

$$\frac{db_{c,t}}{dr_{c,t}} = -a_c\, r^{-\zeta-1}_{c,t}\, b^{1-\zeta}_{c,t} = -\frac{(b^\zeta_{c,t} - 1)}{r_{c,t}}\, b^{1-\zeta}_{c,t},$$

where the last equality uses $a_c\, r^{-\zeta}_{c,t} = b^\zeta_{c,t} - 1$ (from $b^\zeta_{c,t} = 1 + a_c\, r^{-\zeta}_{c,t}$). Therefore:

$$\frac{db_{c,t}}{dd_{c,t}} = \frac{db_{c,t}}{dr_{c,t}} \cdot \frac{dr_{c,t}}{dd_{c,t}} = -\frac{(b^\zeta_{c,t} - 1)\, b^{1-\zeta}_{c,t}}{1 + d_{c,t}}. \tag{C-17}$$

Lastly, from $\tilde{b}_{c,t} = b^{\gamma_c \phi}_{c,t}$, we have

$$\frac{d\tilde{b}_{c,t}}{db_{c,t}} = \gamma_c \phi\, b^{\gamma_c\phi - 1}_{c,t} \frac{db_{c,t}}{dd_{c,t}} = -\frac{\gamma_c \phi\, \tilde{b}_{c,t}}{1 + d_{c,t}} \cdot \frac{b^\zeta_{c,t} - 1}{b^\zeta_{c,t}} = -\frac{\phi\, \gamma_c\, \alpha_{c,t}\, \tilde{b}_{c,t}}{1 + d_{c,t}}, \tag{C-18}$$

where the last equality uses $\frac{b^\zeta_{c,t} - 1}{b^\zeta_{c,t}} = \frac{a_c\, r^{-\zeta}_{c,t}}{1 + a_c\, r^{-\zeta}_{c,t}} = \alpha_{c,t}$. □

Proof of Proposition 4. Since $d\tilde{b}_{k,t}/dd_{c,t} = 0$ for $k \neq c$, from the chain rule we have

$$\frac{d\mathbb{E}[MC_{j,t}]}{dd_{c,t}} = \frac{\partial \mathbb{E}[MC_{j,t}]}{\partial \tilde{b}_{c,t}} \cdot \frac{d\tilde{b}_{c,t}}{dd_{c,t}}.$$

From the expected marginal cost expression, under $F(\varepsilon^k_j) = \varepsilon^k_j / \bar{\varepsilon}$:

$$\mathbb{E}[MC_{j,t}] = c^0_{j,t} - \frac{c^f_{j,t}}{\bar{\varepsilon}} \sum\nolimits_{k=1}^{N} (\varepsilon^k_{j,t})^2, \tag{C-19}$$

where $c^0_{j,t} = C^*_t / \Omega_{j,t}$ and $c^f_{j,t} = C_{F,t}/Y_{j,t}$. Using $\varepsilon^k_{j,t} = (\tilde{b}_{k,t} - 1)\, c^0_{j,t} / (\tilde{B}_{k,t}\, c^f_{j,t})$, this can

be rewritten as:

$$\mathbb{E}[MC_{j,t}] = c^0_{j,t}\left[1 - \frac{1}{\bar{\varepsilon}}\sum\nolimits_{k=1}^{N}\frac{(\tilde{b}_{k,t}-1)}{\tilde{B}_{k,t}}\,\varepsilon^k_{j,t}\right]. \tag{C-20}$$

We differentiate (C-19) term by term. Since $c^0_{j,t}$ and $c^f_{j,t}$ do not depend on $\tilde{b}_{c,t}$ (they involve only aggregate cost indices and firm productivity), the derivative falls entirely on the sum of squared cutoffs:

$$\frac{\partial\,\mathbb{E}[MC_{j,t}]}{\partial\tilde{b}_{c,t}} = -\frac{c^f_{j,t}}{\bar{\varepsilon}}\sum\nolimits_{k=1}^{N}\frac{d\left[(\varepsilon^k_{j,t})^2\right]}{d\tilde{b}_{c,t}} = -\frac{2\,c^f_{j,t}}{\bar{\varepsilon}}\sum\nolimits_{k=1}^{N}\varepsilon^k_{j,t}\frac{d\varepsilon^k_{j,t}}{d\tilde{b}_{c,t}}. \tag{C-21}$$

From $\varepsilon^k_{j,t} = (\tilde{b}_{k,t}-1)\,c^0_{j,t}/(\tilde{B}_{k,t}\,c^f_{j,t})$, there are three cases:

- *For $k < c$:* Neither $\tilde{b}_{k,t}$ nor $\tilde{B}_{k,t}$ involves $\tilde{b}_{c,t}$, so $d\varepsilon^k_{j,t}/d\tilde{b}_{c,t} = 0$.
- *For $k = c$:* Using $\tilde{B}_{c,t} = \tilde{B}_{c-1,t}\,\tilde{b}_{c,t}$ and differentiating $(\tilde{b}_{c,t}-1)/(\tilde{B}_{c-1,t}\,\tilde{b}_{c,t})$ with respect to $\tilde{b}_{c,t}$:
$$\frac{d\varepsilon^c_{j,t}}{d\tilde{b}_{c,t}} = \frac{c^0_{j,t}}{c^f_{j,t}\,\tilde{B}_{c-1,t}\,\tilde{b}^2_{c,t}}. \tag{C-22}$$
- *For $k > c$:* $\tilde{b}_{k,t}$ is independent of $\tilde{b}_{c,t}$, but $\tilde{B}_{k,t} = \tilde{B}_{c-1,t}\,\tilde{b}_{c,t}\prod_{s=c+1}^{k}\tilde{b}_{s,t}$ depends on it through its product structure, giving $d\tilde{B}_{k,t}/d\tilde{b}_{c,t} = \tilde{B}_{k,t}/\tilde{b}_{c,t}$. Since $\varepsilon^k_{j,t} \propto 1/\tilde{B}_{k,t}$:
$$\frac{d\varepsilon^k_{j,t}}{d\tilde{b}_{c,t}} = -\frac{\varepsilon^k_{j,t}}{\tilde{b}_{c,t}}. \tag{C-23}$$

Substituting the above results into (C-21), and noting that the $k < c$ terms vanish:

$$\frac{\partial\mathbb{E}[MC_{j,t}]}{\partial\tilde{b}_{c,t}} = -\frac{2\,c^f_{j,t}}{\bar{\varepsilon}}\left[\varepsilon^c_{j,t}\cdot\frac{c^0_{j,t}}{c^f_{j,t}\,\tilde{B}_{c-1,t}\,\tilde{b}^2_{c,t}} + \sum_{k=c+1}^{N}\varepsilon^k_{j,t}\cdot\left(-\frac{\varepsilon^k_{j,t}}{\tilde{b}_{c,t}}\right)\right]. \tag{C-24}$$

Therefore, we have

$$\frac{d\mathbb{E}[MC_{j,t}]}{dd_{c,t}} = \frac{\partial\mathbb{E}[MC_{j,t}]}{\partial\tilde{b}_{c,t}}\cdot\frac{d\tilde{b}_{c,t}}{dd_{c,t}} = \frac{2\,\phi\,\gamma_c\,\alpha_{c,t}}{\bar{\varepsilon}\,(1+d_{c,t})}\left[c^0_{j,t}\frac{\varepsilon^c_{j,t}}{\tilde{B}_{c,t}} - c^f_{j,t}\sum_{k=c+1}^{N}\left(\varepsilon^k_{j,t}\right)^2\right].$$

Lastly, we show that the elasticity of expected marginal cost with respect to the tariff factor on country c is unambiguously positive: a tariff increase on any supplier

raises expected marginal cost. In fact, we have shown so far that the elasticity of expected marginal cost with respect to the tariff factor on imports from country c is proportional to

$$c_{j,t}^{0}\,\frac{\varepsilon_{j,t}^{c}}{\tilde{B}_{c,t}}-c_{j,t}^{f}\sum_{k=c+1}^{N}\left(\varepsilon_{j,t}^{k}\right)^{2}. \tag{C-25}$$

Factor out $c_{j,t}^{f}$ from (C-25) to write

$$c_{j,t}^{0}\,\frac{\varepsilon_{j,t}^{c}}{\tilde{B}_{c,t}}-c_{j,t}^{f}\sum_{k=c+1}^{N}\left(\varepsilon_{j,t}^{k}\right)^{2}=c_{j,t}^{f}\left[\frac{c_{j,t}^{0}}{c_{j,t}^{f}}\,\frac{\varepsilon_{j,t}^{c}}{\tilde{B}_{c,t}}-\sum_{k=c+1}^{N}\left(\varepsilon_{j,t}^{k}\right)^{2}\right]. \tag{C-26}$$

Next, observe that

$$\frac{TVC_{j,t}^{c}}{C_{F,t}}=\frac{c_{j,t}^{0}}{c_{j,t}^{f}}\,\frac{1}{\tilde{B}_{c,t}}, \tag{C-27}$$

and, by telescoping the cutoff definition (4),

$$\frac{TVC_{j,t}^{c}}{C_{F,t}}=\frac{TVC_{j,t}^{N}}{C_{F,t}}+\sum_{k=c+1}^{N}\varepsilon_{j,t}^{k}. \tag{C-28}$$

Substituting (C-27) and (C-28) into the bracketed term in (C-26) gives

$$\begin{aligned}
&c_{j,t}^{0}\,\frac{\varepsilon_{j,t}^{c}}{\tilde{B}_{c,t}}-c_{j,t}^{f}\sum_{k=c+1}^{N}\left(\varepsilon_{j,t}^{k}\right)^{2}\\
&=c_{j,t}^{f}\left[\varepsilon_{j,t}^{c}\left(\frac{TVC_{j,t}^{N}}{C_{F,t}}+\sum_{k=c+1}^{N}\varepsilon_{j,t}^{k}\right)-\sum_{k=c+1}^{N}\left(\varepsilon_{j,t}^{k}\right)^{2}\right]\\
&=c_{j,t}^{f}\left[\varepsilon_{j,t}^{c}\,\frac{TVC_{j,t}^{N}}{C_{F,t}}+\sum_{k=c+1}^{N}\varepsilon_{j,t}^{k}\left(\varepsilon_{j,t}^{c}-\varepsilon_{j,t}^{k}\right)\right]\\
&>0,
\end{aligned} \tag{C-29}$$

where the strict inequality follows from two observations: (i) $TVC_{j,t}^{N}/C_{F,t}>0$, so the first term is strictly positive; and (ii) the cutoffs are strictly decreasing in rank, so $\varepsilon_{j,t}^{c}>\varepsilon_{j,t}^{k}$ for all $k>c$, which implies that every summand in the second term is strictly positive.

□

Derivation of Equation (33). From (19), we write $s_{j,c,t}$ as a ratio: $s_{j,c,t} = \frac{N_{c,t}}{D_t}$, where we define: $N_{c,t} \equiv \gamma_c \, \alpha_{c,t} \, \mathbf{1}_{jc,t}$, and $D_t \equiv \sum_{c'=1}^{N} \gamma_{c'} \, \alpha_{c',t} \, \mathbf{1}_{jc',t}$. At the steady state, we denote $\bar{N}_c = \gamma_c \, \bar{\alpha}_c \, \bar{\mathbf{1}}_{j,c}$, $\bar{D} = \sum_{c'} \gamma_{c'} \, \bar{\alpha}_{c'} \, \bar{\mathbf{1}}_{jc'}$, and $\bar{s}_{j,c} = \bar{N}_c / \bar{D} = s_{j,c}^{2017}$.

Using the quotient rule $d(N/D) = (dN \cdot D - N \cdot dD)/D^2$, we have

$$\Delta s_{j,c,t} \equiv s_{j,c,t} - \bar{s}_{j,c} \approx \frac{1}{\bar{D}} \left(\Delta N_{c,t} - \bar{s}_{j,c} \, \Delta D_t \right), \tag{C-30}$$

where $\Delta N_{c,t} = N_{c,t} - \bar{N}_c$ and $\Delta D_t = D_t - \bar{D}$

Since $N_{c,t} = \gamma_c \, \alpha_{c,t} \, \mathbf{1}_{j,c,t}$ is a product of two endogenous objects ($\alpha_{c,t}$ and $\mathbf{1}_{j,c,t}$), its first-order change is:

$$\Delta N_{c,t} \approx \gamma_c \, \bar{\mathbf{1}}_{j,c} \, \Delta \alpha_{c,t} + \gamma_c \, \bar{\alpha}_c \, \Delta \mathbf{1}_{j,c,t}. \tag{C-31}$$

Similarly:

$$\Delta D_t \approx \sum_{c'=1}^{N} \left(\gamma_{c'} \, \bar{\mathbf{1}}_{jc'} \, \Delta \alpha_{c',t} + \gamma_{c'} \, \bar{\alpha}_{c'} \, \Delta \mathbf{1}_{j,c',t} \right). \tag{C-32}$$

Substituting (C-31) and (C-32) into (C-30) and separating terms involving country c from those involving $c' \neq c$:

Terms involving $\Delta \alpha_{c,t}$ only (intensive margin for country c):

$$\frac{1}{\bar{D}} \left(\gamma_c \, \bar{\mathbf{1}}_{j,c} \, \Delta \alpha_{c,t} - \bar{s}_{j,c} \, \gamma_c \, \bar{\mathbf{1}}_{j,c} \, \Delta \alpha_{c,t} \right) = \frac{\gamma_c \, \bar{\mathbf{1}}_{j,c}}{\bar{D}} \left(1 - \bar{s}_{j,c} \right) \Delta \alpha_{c,t}.$$

Using $\bar{s}_{j,c} = \gamma_c \, \bar{\alpha}_c \, \bar{\mathbf{1}}_{j,c} / \bar{D}$, this simplifies to:

$$\frac{\bar{s}_{j,c}}{\bar{\alpha}_c} \left(1 - \bar{s}_{j,c} \right) \Delta \tilde{a}_{c,t} \equiv \frac{\partial s_{j,c}}{\partial \alpha_{c,t}} \cdot \Delta \alpha_{c,t}. \tag{C-33}$$

Terms involving $\Delta \mathbf{1}_{jcc,t}$ only (extensive margin for country c):

$$\frac{1}{\bar{D}} \left(\gamma_c \, \bar{\alpha}_c \, \Delta \mathbf{1}_{j,c,t} - \bar{s}_{j,c} \, \gamma_c \, \bar{\alpha}_c \, \Delta \mathbf{1}_{j,c,t} \right) = \frac{\gamma_c \, \bar{\alpha}_c}{\bar{D}} \left(1 - \bar{s}_{j,c} \right) \Delta \mathbf{1}_{j,c,t}.$$

Again using $\bar{s}_{j,c} = \gamma_c \, \bar{\alpha}_c \, \bar{\mathbf{1}}_{j,c} / \bar{D}$:

$$\frac{\bar{s}_{j,c}}{\bar{\mathbf{1}}_{j,c}} \left(1 - \bar{s}_{j,c} \right) \Delta \mathbf{1}_{j,c,t} \equiv \frac{\partial s_{j,c}}{\partial \mathbf{1}_{j,c,t}} \cdot \Delta \mathbf{1}_{j,c,t}. \tag{C-34}$$

Terms involving $\Delta\alpha_{c',t}$ *and* $\Delta\mathbf{1}_{j,c',t}$ *for* $c' \neq c$ (cross-country rebalancing):

$$-\frac{\bar{s}_{j,c}}{\bar{D}}\sum_{c'\neq c}\left(\gamma_{c'}\,\bar{\mathbf{1}}_{j,c'}\,\Delta\alpha_{c',t}+\gamma_{c'}\,\bar{\alpha}_{c'}\,\Delta\mathbf{1}_{j,c',t}\right). \tag{C-35}$$

These terms capture how changes in the expenditure shares or importing status of *other* countries affect $s_{j,c,t}$ through the denominator D_t: if the firm increases imports from country $c' \neq c$ (either by spending more or by entering a new source market), the share of country c mechanically falls.

Combining (C-33), (C-34), and (C-35) yields (33). □

D A Two-Country Simple Case: China and RoW

In this appendix, we consider a simple case with two supplier countries: China and RoW, as studied in Section 5.1. We provide detailed derivations for this simple case.

The movements in the cutoffs classify firms into four types:

- Type A (prob. π_A): *stable single-source* firms with $\varepsilon^{2,post} < \varepsilon_j \leq \varepsilon^{1,post}$. These import from China only in both periods. No change in import status.

- Type B (prob. π_B): *stable dual-source* firms with $\varepsilon_j \leq \varepsilon^{2,pre}$. These import from both countries in both periods (since $\varepsilon^{2,post} > \varepsilon^{2,pre}$, these firms certainly remain dual-sourcers post-war). Their China share adjusts via the *intensive margin* only: $s^{pre} \to s^{post}$ with $s^{post} < s^{pre}$.

- Type C (prob. π_C): *exit importers* with $\varepsilon^{1,post} < \varepsilon_j \leq \varepsilon^{1,pre}$. These were single-source (China only) pre-war but *exit importing entirely* post-war because ε^1 fell below their fixed cost draw. This is the *extensive margin* in action.

- Type D (prob. π_D): *New dual-sourcers*: firms with $\varepsilon^{2,pre} < \varepsilon_j \leq \varepsilon^{2,post}$. These were single-source (China only) pre-war but the rightward shift of ε^2 now makes it optimal to *add RoW* as a second supplier. This is also an extensive margin—but it goes in the direction of *more* importing, not less.

Note that for type A firms, the import share from China remains unchanged at $s_{j,1}^{pre} = s_{j,1}^{post} = 1$, while for type B firms, the higher tariff on China leads them to

import less from China in the post-war period, all else equal:

$$\begin{aligned} pre \quad &: \quad s_{j,1}^{pre} = \frac{\gamma_1 \alpha_1^{pre}}{\gamma_1 \alpha_1^{pre} + \gamma_2 \alpha_2^{pre}} \equiv s^{pre}, \\ post \quad &: \quad s_{j,1}^{post} = \frac{\gamma_1 \alpha_1^{post}}{\gamma_1 \alpha_1^{post} + \gamma_2 \alpha_2^{post}} \equiv s^{post}. \end{aligned}$$

In fact, let $r_{c,t} \equiv \frac{P_{c,t}^M}{V_{c,t}^*}$ denote the relative price of the foreign variety. Log-linearizing around the steady state yields:

$$\widehat{r}_{c,t} = \log r_{c,t} - \log \overline{r}_c = \underbrace{\widehat{e}_{c,t} + \widehat{U}_{c,t} - \widehat{V^*}_{c,t}}_{u_{c,t}} + \Delta \log \left(1 + d_{c,t}\right).$$

Furthermore, given the log-linearization results that $\hat{\alpha}_c = -\zeta \left(1 - \bar{\alpha}_c\right) \widehat{r}_{c,t}$ and $\widehat{b}_{c,t} = -\bar{\alpha}_c \widehat{r}_{c,t}$, we can show $\widehat{s}_1 = \left(1 - \overline{s}_1\right)\left(\hat{\alpha}_1 - \hat{\alpha}_2\right)$, which implies

$$\begin{aligned} s^{post} - s^{pre} &= s^{pre}\left(1 - s^{pre}\right)\left(\widehat{\alpha}_1 - \hat{\alpha}_2\right) \\ &= -\zeta \left(1 - \bar{\alpha}_1\right) s^{pre} \left(1 - s^{pre}\right) \Delta \log \left(1 + d_{1,t}\right) \\ &\quad + \zeta s^{pre}\left(1 - s^{pre}\right)\left(-\left(1 - \overline{\alpha}_1\right) u_{1,t} + \left(1 - \bar{\alpha}_2\right) u_{2,t}\right). \end{aligned}$$

It follows that, all else equal (i.e., fixing the aggregate shocks), an increase in China's tariff rate decreases the import share for type B firms (i.e., $s^{post} < s^{pre}$).

The following table summarizes the results for these four types of firms.

Table D-1: Firm Types

	Type A (stable, China only) prob. π_A	Type B (stable, both) prob. π_B	Type C (exit importing) prob. π_C	Type D (add RoW) prob. π_D
$s_{j,1}^{pre}$	1	s^{pre}	1	1
$s_{j,1}^{post}$	1	s^{post}	–	s^{post}
$\Delta\tau_j^{IV}$	d	$s^{pre} \cdot d$	d	d
$\Delta\tau_j^{OLS}$	d	$s^{post} \cdot d$	0	$s^{post} \cdot d$

Next, we compute the price change for each firm type.

- Type A: unchanged from before—still importing from China at the (now higher)

tariff rate:

$$\Delta p_A = \tilde{\alpha}_1(d + u_1).$$

- Type B: unchanged—still dual-sourcing:

$$\Delta p_B = \tilde{\alpha}_1(d + u_1) + \tilde{\alpha}_2 u_2.$$

- Type C: the firm exits importing. Its pre-war marginal cost was $c_j^0/\widetilde{b}_1^{pre}$ (importing from China), and its post-war marginal cost is c_j^0 (domestic sourcing only). Therefore:

$$\begin{aligned} \Delta p_C &= \log MC_j^{post} - \log MC_j^{pre} \\ &= \log c_j^0 - \log\left(c_j^0/\widetilde{b}_1^{pre}\right) \\ &= \log \widetilde{b}_1^{pre} \equiv \delta_{ext} > 0. \end{aligned}$$

  The term $\delta_{ext} = \log \widetilde{b}_1^{pre} = \gamma_1 \phi \log b_1^{pre} > 0$ is the discrete cost increase from losing the China import variety premium. Importantly, δ_{ext} does not depend on the tariff rate d: the firm's post-war cost is simply the non-importer cost c_j^0, regardless of how large the tariff is, because the firm is no longer importing.[10]

- Type D: the firms that are China-only pre-war ($MC^{pre} = c_j^0/\tilde{b}_1^{pre}$) and dual-source post-war ($MC^{post} = c_j^0/(\tilde{b}_1^{post} \cdot \tilde{b}_2^{post})$). Their price change is:

$$\begin{aligned} \Delta p_D &= \log MC^{post} - \log MC^{pre} \\ &= -(\log \tilde{b}_1^{post} - \log \tilde{b}_1^{pre}) - \log \tilde{b}_2^{post} \\ &= \tilde{\alpha}_1(d + u_1) - \log \tilde{b}_2^{post}. \end{aligned}$$

  The first term is the China tariff pass-through (same as type A). The second term, $-\log \tilde{b}_2^{post} < 0$, is a *cost reduction* from adding RoW—the firm now benefits from the import variety premium of the second country. So type D firms have a *smaller* price increase than type A (they partially offset the China tariff through diversification), yet OLS assigns them *lower* tariff exposure.

[10] More precisely, $\delta_{ext} = \gamma_1 \phi \log b_1^{pre}$ where $b_1^{pre} = [1 + a_1(P_1^{M,pre}/V_1^*)^{-\zeta}]^{1/\zeta}$ reflects the pre-war quality-adjusted cost advantage of Chinese varieties. This is a first-order cost: it equals the full import variety premium that the firm forfeits by exiting.

Let D_j^B, D_j^C and D_j^D be indicators for types B, C and D. To simplify notation, define:

$$\tilde{\alpha}_1 \equiv \bar{\alpha}_1\gamma_1\phi, \qquad \tilde{\alpha}_2 \equiv \bar{\alpha}_2\gamma_2\phi, \qquad d \equiv \Delta\log(1+d_1).$$

We can show that the three key objects can be characterized as follows (see Appendix C for details):

$$\begin{aligned}
\Delta p_j &= \tilde{\alpha}_1(d+u_1)(1-D_j^C) + D_j^B\tilde{\alpha}_2 u_2 + D_j^C\delta_{ext} - D_j^D\log\tilde{b}_2^{post}, \qquad \text{(D-1)}\\
\Delta\tau_j^{OLS} &= d\left[1-(D_j^B+D_j^D)(1-s^{post})-D_j^C\right],\\
\Delta\tau_j^{IV} &= d\left[1-D_j^B(1-s^{pre})\right].
\end{aligned}$$

Note that types A and C have the same $\Delta\tau^{IV} = d$, because both were China-only importers pre-war. The decomposition of the OLS regressor is:

$$\Delta\tau_j^{OLS} = \Delta\tau_j^{IV} + \underbrace{D_j^B(s^{post}-s^{pre})d}_{B_j^{int}} + \underbrace{-D_j^C\cdot d}_{B_j^{ext,C}} + \underbrace{D_j^D(s^{post}-1)d}_{B_j^{ext,D}}. \qquad \text{(D-2)}$$

The intensive-margin compression $B_j^{int} = (s^{post}-s^{pre})d < 0$ affects only type B firms (whose China share falls from s^{pre} to s^{post}). The extensive-margin compression $B_j^{ext,C} = -d < 0$ affects only type C firms (whose OLS regressor drops from d to zero). Crucially, $|B_j^{ext,C}| = d > |B_j^{int}| = (s^{pre}-s^{post})d$: the extensive margin produces a *full* compression (from d to 0), which is much larger than the intensive margin's partial compression (from $s^{pre}d$ to $s^{post}d$). The last term $B_j^{ext,D} = (s^{post}-1)d < 0$ is the extensive-margin compression for type D firms: their OLS regressor is compressed from d to $s^{post}\cdot d$ because they added a new supplier, diluting their measured China exposure.

From the OLS formula: $\hat{\beta}^{OLS} = \frac{\text{Cov}(\Delta p_j, \Delta\tau^{OLS})}{\text{Var}(\Delta\tau^{OLS})}$ and the true structural model (35), we have

$$\hat{\beta}^{OLS} = \underbrace{\beta^{IV}\cdot\frac{\text{Var}(\Delta\tau^{IV})+\text{Cov}(\Delta\tau^{IV},B_j)}{\text{Var}(\Delta\tau^{OLS})}}_{\text{attenuation + substitution bias}} + \underbrace{\frac{\text{Cov}(u_j,B_j)}{\text{Var}(\Delta\tau^{OLS})}}_{\text{endogeneity bias}}, \qquad \text{(D-3)}$$

where $B_j \equiv B_j^{int} + B_j^{ext,C} + B_j^{ext,D}$.

Note that type D firms represent a specific transition (China-only to dual-source) that requires the fixed cost draw to fall in a narrow range $[\varepsilon^{2,pre}, \varepsilon^{2,post}]$. This population is typically small relative to other types. To simplify our discussion, we drop

type D firms.

We now turn to the general case with types A, B and C. For example, if China's tariff hike is large enough, its cutoff moves significantly, leading to a large population of type C firms.

We can now evaluate the endogeneity bias term in equation (D-3). The bias term $B_j^{ext,C} = -d$ is large and negative (type C firms' OLS regressor is compressed to zero), and their structural residual $u_j^{ext,C} = \delta_{ext} - \beta d$ is large and positive (they suffered the discrete variety loss δ_{ext}, which OLS cannot attribute to the tariff because it sees zero tariff exposure for these firms). This makes $\text{Cov}(B_j^{ext,C}, u_j^{ext,C}) < 0$: firms with the most negative bias correction (type C) have the most positive residual. Since the covariance enters the OLS coefficient through $\text{Cov}(\Delta\tau^{OLS}, u) = \text{Cov}(B, u) < 0$, this biases OLS *downward*.

The Bartik IV is immune to this bias. Because the IV assigns $\Delta\tau^{IV} = d$ to *both* type A and type C, it does not condition on the post-war importing decision: whether a firm stays or exits, the IV records the same pre-determined tariff exposure d. The variation in ε_j between types A and C—which is the sole source of the OLS endogeneity—is invisible to the IV because both types receive the same instrument value regardless of their post-war importing status. Formally, $\text{Cov}(\varepsilon_j, \Delta\tau^{IV}) = 0$ within the A+C group (since $\Delta\tau^{IV} = d$ is constant), so the omitted variable bias vanishes under IV estimation.