Estimating Tax-Dependent Compliance: Theory and Evidence from Trade Wars

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Abstract

Statutory tariff rates may overstate the tariffs actually paid due to evasion and avoidance. We develop a novel method to estimate tariff compliance and apply it to the 2018 trade war, when several countries imposed retaliatory tariffs on U.S. exports. Estimated compliance falls by 25 percentage points after tariff increases; a one percentage point tariff increase reduces compliance by 1 to 2 percentage points. Compliance is lower for intermediate goods, which often qualify for duty-free treatment, and for differentiated products, whose valuation is more difficult to verify. The decline accounted for approximately $3.5 billion in foregone tariff revenue in 2019.

Keywords: tariffs, taxation, tax compliance, tax evasion, avoidance, trade war

JEL Classifications: F13, H20, H22, H26, L10

1 Introduction

The 2018 trade war between the United States and its major trading partners marked a sharp departure from decades of trade liberalization. In retaliation against U.S. tariffs, several countries imposed retaliatory tariffs on U.S. exports, with statutory rates rising from an average of 6 percent to 20 percent on affected products. A growing literature documents the effects of these tariffs on trade flows, prices, and welfare (Amiti et al., 2019; Fajgelbaum et al., 2020; Cavallo et al., 2021). This literature typically assumes full compliance with statutory tariff rates. Yet the tariffs actually paid may differ substantially from statutory rates due to both illegal evasion and legal avoidance through duty-drawback provisions, bonded warehouses, free trade zones, and processing trade regimes.

The scope for tariff evasion and avoidance increased sharply on both sides of the 2018 trade war. Evidence from the U.S. import side, where administrative data are more accessible, illustrates the scale: U.S. Customs and Border Protection enforcement actions rose from \$15 million in 2018 to \$215 million in 2020 (Roll and Akers, 2023), and the European Central Bank estimates that U.S. foreign trade zones may have lowered the effective U.S. tariff rate on imports from China by up to 0.7 percentage points, with reductions up to 4.5 percentage points if all affected Chinese intermediates were rerouted through FTZs (Di Nino et al., 2020). Comparable channels operate on the retaliating side: China expanded administrative flexibility in its processing-trade regime in 2019 as retaliatory tariffs on U.S. exports reached their peak (Chor and Li, 2024). These mechanisms have only grown more relevant with the renewed escalation of U.S. tariffs in 2025.[1] Statutory rates may therefore substantially overstate the tariffs actually paid, yet comprehensive estimates of compliance remain elusive, making it difficult to assess the revenue and protective effects of trade policy.

This paper develops a novel method to estimate compliance rates, defined as the fraction

[1] Gopinath and Neiman (2026) document that actual tariff rates on U.S. imports in 2025 were approximately half of statutory rates, reflecting shipping lags, exemptions, USMCA utilization, and enforcement gaps. Recent estimates suggest importers avoided \$6.5 billion in tariffs (13 percent of new revenue) through accelerated purchases and supply chain adjustments during the first months of 2025 (Penn Wharton Budget Model, 2025).

of statutory tariffs reflected in equilibrium outcomes, and applies it to U.S. exports during the 2018 trade war. The standard approach following Fisman and Wei (2004) infers evasion from gaps between importer and exporter reports of the same trade flow. This method identifies how evasion *changes* with tariff rates but cannot recover the *level* of compliance. Our approach fills this gap: we estimate both the level of compliance and its variation with tariff rates, relying only on standard trade data widely available across countries and time periods, as well as publicly available input-output accounts.

Our identification strategy builds on the framework outlined by Zoutman et al. (2018) and Bibler et al. (forthcoming). In our setting, foreign importers bear the statutory burden of tariffs on U.S. exports. The empirical challenge is to separately identify the demand elasticity and the compliance rate governing how statutory tariffs map into effective trade costs. Intuitively, the magnitude of the demand response to tariff changes depends on both the elasticity of demand and the compliance rate, so variation in tariff rates alone is not sufficient to separately identify the two parameters. We address this by contrasting responses to (i) supply-side cost shifters and (ii) changes in the tariff rate. A supply shifter that affects U.S. production costs but not foreign demand traces out the demand curve, identifying the demand elasticity. Given the demand elasticity, the response of quantities to tariff changes identifies compliance: under full compliance, quantities fall with the elasticity-implied magnitude, while attenuated responses indicate incomplete implementation. We extend the framework by estimating a flexible compliance function that allows compliance to vary with the tariff level. Although we focus on demand-side statutory incidence, the identification strategy outlined in our framework adapts seamlessly when the statutory burden falls on sellers. Our approach to identifying compliance is also robust to standard forms of imperfect competition.

We implement this strategy using detailed monthly U.S. export data from January 2017 to January 2020, covering the universe of U.S. exports to six retaliating economies: China, Canada, Mexico, the European Union, Turkey, and Russia. We construct applied tariff rates that account for preferential trade agreements and address potential endogeneity from strategic tariff targeting following Boehm et al. (2023).

We begin with an event-study analysis to document the reduced-form effects of retaliatory tariffs. Consistent with Amiti et al. (2019) and Fajgelbaum et al. (2020), we find no evidence of anticipatory responses, sharp declines in export quantities of 15–25 percent within six months of tariff imposition, and no significant change in exporter prices. These patterns motivate our structural approach: the absence of exporter price adjustment is consistent with tariff incidence falling primarily on importers, while the magnitude of the quantity response reflects a combination of demand elasticity and compliance that cannot be separately identified from the reduced-form evidence alone.

For the supply shifter, we adopt a Bartik shift-share design that interacts predetermined industry-level fossil fuel cost shares (from the 2016 input-output tables) with monthly variation in fossil fuel prices. Conditional on importer country-month fixed effects, differential exposure to energy cost shocks generates plausibly exogenous variation in export supply across products. We parameterize the compliance function as a flexible polynomial in the tariff rate, with the polynomial order selected via cross-validation.

We estimate a demand elasticity of -1.7 for U.S. exports, broadly consistent with short-run trade elasticities in the literature (Amiti et al., 2019; Fajgelbaum et al., 2020; Boehm et al., 2023). The estimated compliance function implies that effective compliance was approximately 78 percent at pre-war tariff levels and fell to 53 percent after the tariff increases, a decline of approximately 25 percentage points. The marginal effect of a tariff increase on compliance is negative: a one percentage point increase in the tariff rate reduces compliance by 1 to 2 percentage points, suggesting that higher tariffs are partially self-defeating. As an internal validity check, the estimated compliance function implies near-full compliance when tariffs approach zero, consistent with the prior that there is little to gain from evasion or avoidance when the tariff burden is negligible. These estimates should be interpreted as short-run compliance, reflecting evasion and legal avoidance on continuing trade flows within the adjustment window captured by our data. Longer-run responses, including trade diversion and production relocation to third countries, would attenuate the quantity response through a distinct margin and are outside the scope of our framework. Our findings are robust to alternative instrument constructions and fixed effects specifications, including demanding specifications with industry-month-year fixed effects that absorb sector-specific demand shocks.

Three results merit emphasis. First, the estimated compliance rate of 78 percent at pre-war tariff levels implies a degree of non-compliance, including in trade between developed economies. We interpret this as reflecting both illegal evasion and legal avoidance mechanisms (duty-drawback provisions, bonded warehouses, free trade zones, and processing trade regimes). Supporting this interpretation, we find that compliance is systematically lower for intermediate goods, which are more likely to qualify for duty-free treatment, than for final goods.

Second, compliance for non-differentiated products is near unity, while compliance for differentiated products is substantially lower. This pattern is consistent with Javorcik and Narciso (2008): homogeneous goods have observable reference prices that customs can verify, whereas differentiated products lack clear benchmarks, creating scope for evasion through misreported quantities and values.

Third, translating our estimates into revenue terms, the decline in compliance during the trade war accounts for approximately \$3.5 billion in foregone tariff revenue for U.S. trading partners in 2019, roughly 36 percent of the total revenue gap that year (with the remainder reflecting non-compliance that predates the tariff increases).

Our findings contribute to three strands of literature. First, we provide the first estimates of the equilibrium *level* of tariff compliance, complementing the literature following Fisman and Wei (2004), which estimates how compliance *changes* with tariff rates. Our marginal effects are comparable to Fisman and Wei (2004), who find that a 1 percentage point increase in the tariff rate increases evasion by 3 percent: expressed as a percent change in evasion, our estimates imply a 3 to 9 percent increase. The range includes their estimate, and the slightly larger upper bound is consistent with our broader measure capturing changes in both legal avoidance and illegal evasion, whereas the trade-gap approach primarily reflects quantity and value discrepancies associated with illegal evasion. Second, we contribute to the literature on the 2018 trade war (Amiti et al., 2019; Fajgelbaum et al., 2020; Cavallo et al., 2021; Boehm et al., 2023) by showing that incomplete compliance attenuates the effective tariff increases faced by U.S. exporters. Third, we extend the literature on estimating tax compliance from tax variation (Bibler et al., 2021, forthcoming) by allowing compliance to depend on the tax level, a theoretically motivated and empirically important generalization.

The remainder of the paper is organized as follows. Section 2 provides background on the 2018 trade war and mechanisms of tariff non-compliance. Section 3 develops the theoretical framework and identification strategy. Section 4 describes the data and presents event-study evidence. Section 5 details the empirical specification. Section 6 presents the main results, robustness checks, and heterogeneity analyses. Section 7 concludes.

2 Background

2.1 The 2018-2019 Trade War

In January 2018, the United States imposed global safeguard tariffs on imports of solar panels and washing machines, marking the beginning of a broader trade conflict. These measures escalated over time, extending to over 10,000 product varieties by the end of 2019, with more than half originating from China (Handley et al., 2025). This escalation prompted significant retaliation from key U.S. trading partners, including China, the European Union (EU), Turkey, Mexico, Canada, and Russia.[2]

We analyze the effects of these retaliatory tariffs imposed on U.S. exports. To do so, we

[2]We exclude the last two retaliatory waves, India (June 2019) and the third round of Chinese tariffs (September 2019), because the short remaining sample period provides insufficient post-treatment variation.

build a monthly panel of U.S. exports and bilateral tariffs faced by U.S. exporters during January 2017–January 2020, as described in Section 4.[3]

Table 1 summarizes the timeline, scope, and variation of the retaliatory tariffs on U.S. exports in 2018–2019. Columns 1–2 document the timing of the trade war, listing countries by the month in which they first implemented retaliatory tariffs. Columns 3–5 show the total scope of affected exports, indicating that retaliations targeted approximately $117 billion in U.S. annual exports, representing 10.7% of total U.S. exports in 2017 and covering 7,175 product varieties. Columns 6–8 report the average tariffs on U.S. exports, which increased substantially during this period. The trade-weighted average tariff increased from 5.9% to 11.1% from 2017 to 2018, and further increased to 20.4% by 2019. The table highlights substantial cross-sectional and temporal variation that we exploit in our empirical application.

Figure 1 illustrates the evolution of retaliatory tariffs from January 2018 to January 2020. Panel (a) presents the average additional tariffs weighted by trade value by country and month, showing that China imposed the highest tariff increases among U.S. trade partners. Retaliatory tariffs largely persisted through 2020 for most countries, except for Canada and Mexico, whose additional tariffs fell to zero following the submission of the United States-Mexico-Canada Agreement to the U.S. Congress on May 30, 2019. Panel (b) shows the number of product varieties subject to retaliation. China targeted the largest number of product varieties, followed by Canada. Seasonal patterns are evident, with January declines reflecting typical seasonality in trade flows.

2.2 Tariff Compliance

A large literature documents that departures from full tariff implementation are economically meaningful across institutional settings, product types, and time periods, including evidence from developed countries (Ferrantino et al., 2012; Stoyanov, 2012), developing and transition

[3] Compared to the data used in Cavallo et al. (2021), we incorporate the exemptions on automobile parts that were extended from Dec 15, 2019 to all later months.

economies (Mishra et al., 2008; Epaphra, 2015; Sequeira, 2016), and broader cross-country samples (Buehn and Eichler, 2011; Kellenberg and Levinson, 2019; Kitsios et al., 2023). Illegal tariff non-compliance occurs through three main mechanisms: (i) underreporting taxable quantities, (ii) underreporting unit values, and (iii) misclassification of products into lower-tax categories (Sequeira, 2016). A robust empirical regularity is that evasion incentives increase with tariff rates, consistent with the seminal evidence in Fisman and Wei (2004).

The prevalence and margin of evasion vary with product characteristics and with the structure of the tariff schedule. Evasion through underreporting of unit values is generally more feasible for differentiated products, where customs valuation is intrinsically more difficult (Javorcik and Narciso, 2008). Misclassification is typically local in the Harmonized System (HS), the standard product coding system used in customs statistics: it is more plausible to reclassify a shipment into a closely related code than into a clearly distinct category. Thus, misclassification often occurs within a broad four-digit HS heading (HS4) by reporting the good under an alternative, more disaggregated six-digit code (HS6) or finer within the same HS4 heading. This incentive is strongest when tariff rates vary substantially across HS6 subheadings within a given HS4 heading (Fisman and Wei, 2004).

Because evasion is unobserved by design, directly measuring it would require combining exporters' true shipments with importers' realized tariff liabilities (or revenues) at the transaction level; such information is rarely available and difficult to assemble across countries, especially since exporters have little incentive to facilitate foreign customs enforcement. As a result, much of the empirical evidence infers evasion from discrepancies between exporter- and importer-reported trade statistics ("trade gaps"), following Fisman and Wei (2004). While these gaps are an imperfect proxy for illegal activity, reflecting not only evasion but also differences in timing of recording and measurement discrepancies in trade statistics, they can be sizable in the aggregate: Kitsios et al. (2023) report crude discrepancies ranging from 2.4 percent of GDP in developed countries to 6.6 percent in low-income countries.

The gap-based approach has been used to study how enforcement and institutions shape compliance. Direct enforcement actions, reductions in administrative frictions, and digitization of customs procedures have been shown to reduce evasion, as documented by Mishra et al. (2008), Beverelli and Ticku (2022), and Kitsios et al. (2023). International rules also matter. The WTO Customs Valuation Agreement (CVA), implementing GATT Article VII, limits discretion in customs valuation by requiring customs authorities in importing countries to accept the invoice values declared by exporters. Using WTO accession episodes, Javorcik and Narciso (2017) find that accession significantly reduces underreporting of import prices. However, changes to customs valuation procedures induce importers to seek alternative evasion methods, such as underreporting quantities and product misclassification, leading to no change in overall evasion.

Observed departures from full tariff implementation need not reflect illegal evasion alone. The effective incidence of statutory tariffs can also be reduced through legal avoidance mechanisms that eliminate or suspend tariff liability for eligible transactions. These mechanisms operate through several channels. Duty-drawback provisions allow importers to claim refunds on tariffs paid on inputs that are subsequently incorporated into exported goods, effectively rendering the tariff on those inputs non-binding for export-oriented firms. Bonded warehouses permit firms to store imported goods without paying duties until the goods enter domestic commerce, enabling deferral or full avoidance if goods are ultimately re-exported. Free trade zones suspend tariff liability within the zone boundary, though the extent of FTZ-based duty relief varies across jurisdictions: U.S. FTZs grant extensive exemptions, Chinese FTZs currently do not waive import duties, and the European Union relies on a separate duty-suspension scheme (Di Nino et al., 2020). Processing-trade regimes, which allow firms to import inputs duty-free conditional on exporting the processed output, perform an analogous function and are particularly prevalent in China. Because these regimes are designed to facilitate the import of inputs into export supply chains, intermediate goods are systematically more likely to qualify for duty-free treatment than final goods (Grant, 2020).

The scale of legal avoidance during the 2018–2019 trade war was quantitatively significant. While our focus is on retaliatory tariffs imposed on U.S. exports, evidence from the U.S. import side highlights the broader magnitude of avoidance behavior. The European Central Bank estimates that recourse to free trade zones may have lowered the effective U.S.-China tariff rate by up to 0.7 percentage points on U.S. imports from China, with potential reductions of up to 4.5 percentage points if all affected intermediates had been rerouted through free trade zones (Di Nino et al., 2020). Consistent with these patterns, China reduced administrative barriers for firms engaging in processing trade beginning January 1, 2019 (Chor and Li, 2024), expanding the scope for legal avoidance on the retaliatory margin precisely as tariffs reached their peak. Taken together, these patterns imply that the effective tariff burden faced by importers likely fell short of statutory rates, particularly for intermediate goods and firms integrated into global supply chains.

Existing evidence on compliance during the trade war focuses primarily on the U.S. import side. Che et al. (2025) document increased tariff evasion on U.S. imports during the trade war, with stronger responses among intermediate goods and among importers facing larger tariff liabilities. A related margin is re-routing through third countries: Iyoha et al. (2024) estimate that re-routing of Chinese exports to the United States via Vietnam accounted for 1.8 percent of Vietnamese exports to the United States in 2021, suggesting that re-routing was modest relative to broader trade diversion. Much less is known about compliance and avoidance on the retaliating side, that is, in the countries imposing tariffs on U.S. exports, largely due to limited access to administrative customs data. This is the focus of our paper.

Taken together, this evidence indicates that statutory tariffs may translate imperfectly into effective tariff burdens through both illegal evasion and legal avoidance. Because these channels operate through the same observed trade aggregates and direct measures of tariff payments are typically unavailable, we treat observed responses to tariff changes as reflecting their combined effect in the empirical analysis below.

3 Identifying Compliance from Tax Rate Variation

3.1 Identification of Tax Compliance

We develop a framework to identify the effective tax compliance, denoted by λ, which measures the fraction of the statutory tax that is effectively implemented in equilibrium outcomes. Our key insight is that by observing how equilibrium prices and quantities respond to changes in taxes and an independent supply shifter, we can separately identify the *demand elasticity* and the *effective compliance rate*.

Our identification strategy is formulated for a generic ad-valorem tax. In the empirical application, this tax corresponds to an import tariff. Our approach extends the models of Romalis (2007) and, in particular, Zoutman et al. (2018), who assume full compliance. We also build on Bibler et al. (forthcoming), differing in that we rely solely on statutory tax-rate variation rather than changes in tax policies.

We observe panel data for a good i at time t, including equilibrium tax-exclusive prices P_{it} and quantities Q_{it}. The logged structural demand and supply equations are:

$$
\begin{aligned}
q_{it} &= \varepsilon^d p_{it} + \gamma^d T_{it} + \rho^d Z_{it} + v_{it}^d, \\
q_{it} &= \varepsilon^s p_{it} + \gamma^s T_{it} + \rho^s Z_{it} + v_{it}^s,
\end{aligned}
$$

where q_{it} and p_{it} denote the logarithms of quantity and price. Let $T_{it} = f(\tau_{it})$ denote a function of the ad-valorem tax rate τ_{it}, and Z_{it} is an additional demand or supply shifter. In our application, we take $T_{it} = \ln(1 + \tau_{it})$, so that the effective tax wedge is $\lambda(T_{it})T_{it}$. As the equations are expressed in logarithms, the price coefficients correspond to the structural demand and supply elasticities.

Equating demand and supply gives the reduced-form relations:

$$
\begin{aligned}
q_{it} &= \pi_{Tq} T_{it} + \pi_{Zq} Z_{it} + \zeta_{it}^{q}, \\
p_{it} &= \pi_{Tp} T_{it} + \pi_{Zp} Z_{it} + \zeta_{it}^{p},
\end{aligned}
$$

where the reduced-form coefficients are functions of the underlying structural parameters:

$$
\begin{aligned}
\pi_{Tq} &= \frac{\gamma^d \varepsilon^s - \gamma^s \varepsilon^d}{\varepsilon^s - \varepsilon^d}, \qquad \pi_{Tp} = \frac{\gamma^d - \gamma^s}{\varepsilon^s - \varepsilon^d}, \\
\pi_{Zq} &= \frac{\rho^d \varepsilon^s - \rho^s \varepsilon^d}{\varepsilon^s - \varepsilon^d}, \qquad \pi_{Zp} = \frac{\rho^d - \rho^s}{\varepsilon^s - \varepsilon^d}.
\end{aligned}
$$

Identification Strategy. We consider the market for U.S. exports, where the tax (an import tariff) is levied on the demand side, so that buyers in the importing country bear the statutory burden. When taxes are levied on the demand side, identification relies on three restrictions:

Assumption 1. ***Standard Exclusion Restriction (SER):*** *The tax is levied on the demand side, implying $\gamma^s = 0$.*

Assumption 2. ***Second Standard Exclusion Restriction (SER2):*** *The variable Z_{it} acts as a pure supply shifter, implying $\rho^d = 0$.*[4]

Assumption 3. ***Ramsey Exclusion Restriction with Partial Compliance (RER):*** *Demand depends (locally) on the price after taxation, and tax variation shifts demand in proportion to both the demand elasticity and the compliance rate:*

$$
\gamma^d = \lambda(T_{it}; \Theta)\varepsilon^d,
$$

where $\lambda(\cdot; \Theta)$ may be constant or depend on the tax level, and Θ is a parameter vector. This restriction rules out direct effects of the tax on demand other than through the tax-inclusive

[4]As outlined by Bibler et al. (forthcoming) and shown in Appendix A.1, whether a supply or demand shifter is needed depends on the statutory burden of the tax.

price.

Under these assumptions, the structural system simplifies to:

$$\begin{aligned} q_{it} &= \varepsilon^d p_{it} + \lambda(T_{it};\Theta)\varepsilon^d T_{it} + v_{it}^d, \\ q_{it} &= \varepsilon^s p_{it} + \rho^s Z_{it} + v_{it}^s. \end{aligned}$$

Equating demand and supply yields the following equilibrium relations:

$$\begin{aligned} q_{it} &= \frac{\lambda(T_{it};\Theta)\varepsilon^d\varepsilon^s}{\varepsilon^s-\varepsilon^d}T_{it} + \frac{-\rho^s\varepsilon^d}{\varepsilon^s-\varepsilon^d}Z_{it} + \zeta_{it}^q, \\ p_{it} &= \frac{\lambda(T_{it};\Theta)\varepsilon^d}{\varepsilon^s-\varepsilon^d}T_{it} + \frac{-\rho^s}{\varepsilon^s-\varepsilon^d}Z_{it} + \zeta_{it}^p. \end{aligned}$$

Identification with Constant Compliance. When the compliance rate does not depend on the tax level, $\lambda(T;\Theta) = \lambda$, and the reduced form equations can be written as:

$$\begin{aligned} q_{it} &= \frac{\lambda\varepsilon^d\varepsilon^s}{\varepsilon^s-\varepsilon^d}T_{it} + \frac{-\rho^s\varepsilon^d}{\varepsilon^s-\varepsilon^d}Z_{it} + \zeta_{it}^q, \\ p_{it} &= \frac{\lambda\varepsilon^d}{\varepsilon^s-\varepsilon^d}T_{it} + \frac{-\rho^s}{\varepsilon^s-\varepsilon^d}Z_{it} + \zeta_{it}^p. \end{aligned}$$

The reduced-form coefficients under constant compliance thus satisfy:

$$\begin{aligned} \pi_{Tq} &= \frac{\lambda\varepsilon^d\varepsilon^s}{\varepsilon^s-\varepsilon^d}, \qquad \pi_{Tp} = \frac{\lambda\varepsilon^d}{\varepsilon^s-\varepsilon^d}, \\ \pi_{Zq} &= \frac{-\rho^s\varepsilon^d}{\varepsilon^s-\varepsilon^d}, \qquad \pi_{Zp} = \frac{-\rho^s}{\varepsilon^s-\varepsilon^d}. \end{aligned}$$

Under Assumptions 1 and 2, we can express the supply and demand elasticities in terms of reduced-form coefficients:

$$\varepsilon^d = \frac{\pi_{Zq}}{\pi_{Zp}}, \tag{1}$$

$$\varepsilon^s = \frac{\pi_{Tq}}{\pi_{Tp}}. \tag{2}$$

In addition, we can solve for γ^d as follows:

$$\gamma^d = \pi_{Tp}(\varepsilon^s - \varepsilon^d). \tag{3}$$

Combining Assumption 3, which simplifies to $\gamma^d = \lambda\varepsilon^d$, with Equations (1), (2), and (3), the compliance rate is identified as:

$$\lambda = -\pi_{Tp} + \frac{\pi_{Tq}}{\varepsilon^d}. \tag{4}$$

In the demand-side incidence case, a higher tax shifts demand inward and reduces the equilibrium tax-exclusive price, implying $\pi_{Tp} < 0$ and the first term in Equation (4) is positive. The compliance rate captures the extent to which the tax-induced demand shift translates into effective price and quantity adjustments.

Considering the case with constant compliance is instructive for two reasons. First, when effective compliance does not vary with the tax rate, the model is fully identified from reduced-form responses to (i) tax changes and (ii) a single supply shifter. In particular, Z identifies ε^d via (1), while T identifies ε^s via (2); given ε^d, (4) then identifies λ. Second, (4) clarifies why compliance is separately identified from the demand elasticity: tax variation delivers price and quantity responses (π_{Tp}, π_{Tq}), while recovering the effective compliance requires ε^d, which is pinned down by independent variation in Z under Assumption 2.

Identification with Tax-Dependent Compliance. Consider the reduced-form partial effects of T and Z, which satisfy:

$$\pi_{Tq}(T) = \frac{\varepsilon^d\varepsilon^s}{\varepsilon^s - \varepsilon^d}\left[\lambda'(T;\Theta)T + \lambda(T;\Theta)\right], \qquad \pi_{Tp}(T) = \frac{\varepsilon^d}{\varepsilon^s - \varepsilon^d}\left[\lambda'(T;\Theta)T + \lambda(T;\Theta)\right],$$

$$\pi_{Zq} = \frac{-\rho^s\varepsilon^d}{\varepsilon^s - \varepsilon^d}, \qquad \pi_{Zp} = \frac{-\rho^s}{\varepsilon^s - \varepsilon^d}.$$

Here, $\pi_{Tq}(T)$ and $\pi_{Tp}(T)$ indicate that the marginal effects in the reduced-form equations are functions of T (and Θ). From the exclusion restrictions, the demand elasticity is identified from the ratio of coefficients on the supply shifter:

$$\varepsilon^d = \frac{\pi_{Zq}}{\pi_{Zp}}.$$

In other words, estimation of the demand elasticity by instrumenting for prices using Z is unaffected by allowing λ to depend on T.

As previously defined, $\pi_{Tp}(T)$ and $\pi_{Tq}(T)$ denote the local reduced-form responses of prices and quantities to the tax. Because the tax term in equilibrium enters as $\lambda(T;\Theta)T$, applying the product rule implies $\partial(\lambda(T;\Theta)T)/\partial T = \lambda'(T;\Theta)T + \lambda(T;\Theta)$. Given ε^d, the reduced-form responses $(\pi_{Tp}(T), \pi_{Tq}(T))$ are therefore sufficient to recover $\lambda'(T;\Theta)T + \lambda(T;\Theta)$:

$$\lambda'(T;\Theta)T + \lambda(T;\Theta) = -\pi_{Tp}(T) + \frac{\pi_{Tq}(T)}{\varepsilon^d}.$$

This equation defines the mapping between observable reduced-form responses and the compliance function $\lambda(\cdot;\Theta)$. Under a parameterization of $\lambda(\cdot;\Theta)$, the shape of $\lambda(T)$ (and, hence, the parameter vector Θ) is identified from the observed pass-through of the tax to prices and quantities. For this relationship to hold, we require only mild regularity conditions on $\lambda(\cdot)$: $\lambda(T)$ must be continuously differentiable on the support of T, ensuring that $\lambda'(T)$ exists and that the differential equation above is well-defined. We therefore allow flexible, continuously differentiable specifications for $\lambda(T)$, such as polynomials or splines. In our empirical implementation, we estimate the structural demand equation using instrumental variables and parameterize $\lambda(T)$ flexibly, rather than estimating the reduced-form coefficients explicitly.

The key insight is that *exogenous variation in* Z_{it} identifies ε^d, while *variation in* T_{it} identifies $\lambda(T)$. Although the supply elasticity ε^s affects the individual reduced-form coefficients π_{Tp} and π_{Tq}, it cancels out in the expression that defines $\lambda(T)$. This allows us to identify the compliance parameter without directly estimating ε^s. This can also be seen by considering estimation of only the demand function; exogenous variation in T and Z is sufficient to separately identify ε^d and $\lambda(T;\Theta)$.

Economically, we expect $\lambda(T)$ to be decreasing in the tax level: as the statutory tax rises, the incentives for evasion and avoidance strengthen, leading to a lower fraction of the tax being effectively remitted. Formally, if $\lambda'(T) < 0$, higher taxes dampen the effective demand shift induced by tax changes, attenuating the observed pass-through of taxes to prices and quantities. In the constant-compliance benchmark, $\lambda'(T) = 0$ and tax changes shift demand proportionally to $\varepsilon^d\lambda$; in contrast, when $\lambda'(T) < 0$, the magnitude of this shift is attenuated relative to the constant-compliance case. This comparative static motivates allowing $\lambda(T)$ to vary flexibly, possibly in a nonlinear fashion, with the tax rate.

Economic Intuition. Figure 2a illustrates the identification of the compliance rate when compliance does not depend on the tax level. A tax increase shifts the demand curve downward by an amount proportional to the effective compliance rate $\lambda\varepsilon^d$, thereby moving the equilibrium along the supply curve. The demand elasticity ε^d is identified by the supply shifter Z, which shifts supply without affecting demand. Given ε^d, the response of equilibrium prices and quantities to tax changes pins down λ via (4). Intuitively, independent variation in Z traces out the demand curve (identifying ε^d), while variation in T shifts demand; combining these two sources of variation disentangles compliance (λ) from the underlying demand elasticity.

Figure 2b illustrates identification when compliance depends on the tax level. A tax increase still shifts demand inward and moves the equilibrium along the supply curve, but the magnitude of the shift now varies with T because the effective wedge is $\lambda(T)T$. If compliance declines as taxes rise, demand shifts become smaller in absolute value, attenuating the pass-through of taxes to prices and quantities. Given ε^d identified from the supply shifter, variation in the joint price and quantity responses to T identifies the tax-dependence of $\lambda(T)$.

While our baseline analysis focuses on demand-side statutory incidence, relevant for trade taxes borne by importers, the estimation strategy adapts seamlessly when the statutory burden falls on the supply side, as shown in Appendix A.1. In that case, a change in the tax rate shifts the *supply* curve proportionally to the constant (or locally varying) compliance rate, and identification proceeds by swapping the roles of the shifters: a *demand* shifter Z_{it} identifies the supply elasticity, while tax variation identifies the demand elasticity. Given ε^s, the compliance rate, λ, is then recovered from the reduced-form responses to the tax.

Finally, Appendix A.2 shows that our compliance estimators are robust to standard forms of imperfect competition. While imperfect competition affects pass-through, changing the relative response of prices and quantities to tax shocks, this scaling appears symmetrically in both reduced-form equations. Thus, the ratios that identify the structural elasticities and the compliance parameter remain invariant.

3.2 Application to Import Tariffs

We apply the identification framework above to import tariffs levied on U.S. exports. The key empirical objects are tax-exclusive (exporter) prices and quantities. In the data, we observe customs values and physical quantities at the product–country–time level rather than transaction-level prices; following standard practice in the trade literature, we therefore interpret unit values constructed from these data as measures of tax-exclusive exporter prices.

Maintaining Assumptions 1–3, supply depends on the tax-exclusive price received by exporters, while demand depends on the tax-inclusive price faced by importers, with statutory tariff variation entering demand through the effective implementation rate. In this setting, departures from full implementation are therefore reflected in a *quantity gap*: relative to the benchmark of full compliance, tariff increases generate attenuated responses of quantities (and tax-exclusive prices) to statutory changes. Our identification strategy recovers this wedge by comparing the responses to tariff reforms with the responses induced by an independent supply shifter. We argue below that the quantity channel is plausibly the empirically relevant margin in this setting, as WTO customs valuation rules constrain price-based manipulation.

Assumption 3, the Ramsey Exclusion Restriction, is standard in the tax incidence literature (Zoutman et al., 2018). The assumption requires that tariffs affect demand only through their effect on tax-inclusive prices, ruling out direct effects of the tariff rate on demand conditional on prices. Violations could arise if tariff changes generate policy uncertainty that independently affects import demand, or if importers interpret tariff increases as signals of future trade policy beyond their direct price effects. In our setting, such concerns are mitigated by the fact that retaliatory tariffs were largely implemented rapidly and publicly announced, reducing the scope for differential uncertainty across varieties. To the extent that any residual uncertainty affected demand uniformly across products within a country, the empirical specification uses importer-month-year fixed effects to absorb it.

Effective implementation may depart from full compliance through several channels. In particular, the fraction of the statutory tariff that is effectively paid can be reduced by both illegal evasion and legal avoidance. As discussed in Section 2.2, the trade literature identifies three mechanisms of tariff evasion: (i) underreporting of unit values, (ii) misclassifying higher-taxed products as lower-taxed varieties, and (iii) underreporting taxable quantities. Quantity gaps arising from misclassification and from underreporting quantities are difficult to distinguish in our data, but both are captured by our framework under the umbrella of the effective compliance rate λ. Legal avoidance can similarly reduce effective compliance through duty-free regimes, exemptions, and related institutional features that eliminate tariff liability for specific transactions. For this reason, we interpret λ as an *effective compliance (implementation) rate*, defined as the fraction of the statutory tariff reflected in observed equilibrium responses under our maintained assumptions, irrespective of whether the underlying mechanism is illegal evasion or legal avoidance.

We argue that the quantity-gap channel is likely to be the empirically relevant margin in many trade environments. Among WTO members, systematic manipulation of customs values is plausibly constrained by the rules governing customs valuation. Under Article VII of the General Agreement on Tariffs and Trade (GATT), customs values are intended to be based on transaction values declared by exporters, which limits discretion in determining dutiable values and reduces scope for price-based manipulation at the border. Consistent with this, Javorcik and Narciso (2017) document that underreporting of unit values declines following WTO accession and that importers substitute toward alternative evasion mechanisms, including misclassification and underreporting of quantities, when tariffs rise. This underscores the relevance of focusing on quantity gaps as a key mechanism of tariff evasion and highlights how our main assumptions likely hold for this type of tariff evasion.

Protectionist measures can incentivize firms to relocate production to jurisdictions not subject to tariffs, a strategy often termed "country hopping" (Flaaen et al., 2020; Utar, 2026). However, the feasibility of this response is highly time-dependent. In the short run, the relevant time horizon for our empirical application, production capacities and entrenched supply chains are relatively inflexible in the aftermath of tariff changes. Consistent with this, recent evidence from the solar panel industry shows that even for geographically diversified firms with existing facilities outside the targeted country (China), relocating production took approximately one to one and a half years after the imposition of tariffs (Bollinger et al., 2024). This substantial lag confirms that tariff avoidance through relocation is primarily a medium- to long-run strategy. Consequently, in the short run, we expect adjustments to be dominated by trade-flow and pricing responses rather than production relocation.

A related concern is that importers may substitute toward third-country suppliers rather than comply with higher tariffs, reducing U.S. export quantities beyond what the tariff-induced price effect alone would predict. If present, this would amplify observed quantity responses to tariffs and bias our compliance estimates upward, making our finding of substantially incomplete post-war compliance conservative. Fajgelbaum et al. (2024) document that bystander countries expanded exports in tariff-exposed products during the trade war, confirming that cross-country reallocation did occur. However, the scope for this channel to affect our estimates is limited. Fajgelbaum et al. (2024) find that bystander countries barely increased their exports to China in response to Chinese retaliatory tariffs on the United States, the destination that constitutes our largest source of tariff variation. To the extent that third-country substitution did occur, it would bias our compliance estimates upward, so our finding of substantially incomplete compliance should be interpreted as a lower bound on true non-compliance.

Finally, our empirical analysis uses multiple goods (or product varieties). In Appendix A of their study, Zoutman et al. (2018) show that the identification results generalize to settings with multiple goods as long as they face independent variation in the tax rates. In our setting, this result implies a straightforward modification of Assumption 3, the Ramsey Exclusion Restriction with Partial Compliance (RER) with multiple goods. Demand for each good depends only on the price after the application of the good-specific tax, and variations in the tax rates induce demand shifts that depend on the degree of compliance and the demand elasticity.

Our analytical framework establishes that exogenous variation in the supply shifter Z_{it} and in the tariff rate T_{it} is sufficient to identify the demand elasticity ε^d and the compliance function $\lambda(T)$. The supply shifter isolates movements along the demand curve, identifying ε^d, while tariff variation shifts demand itself, identifying $\lambda(T)$ conditional on ε^d. Since the supply elasticity ε^s enters only proportionally in both reduced-form equations, it need not be estimated separately. Next, we translate this identification logic into an empirical specification suitable for panel-data estimation of U.S. exports, allowing us to recover the compliance function directly from observed price and quantity responses to tariff changes.

4 Data

We build a monthly panel dataset of applied tariffs on U.S. exports from January 2017 to January 2020. We obtain the retaliatory tariff schedules, including implementation dates and affected Harmonized System (HS) lines, from Cavallo et al. (2021), who compile these measures from official sources catalogued by the International Trade Administration. We measure baseline applied tariffs using implemented rates from the Trade Analysis Information System (TRAINS) database, and construct applied tariffs for each importer–product–month as the sum of the baseline applied rate (measured in 2017) and any additional retaliatory increment imposed after 2017. Tariffs are measured at the six-digit Harmonized System (HS6) level.[5]

Because baseline tariff rates reflect not only *most-favored-nation (MFN)* rates applied to all World Trade Organization members but also *preferential rates* granted under free trade agreements, it is essential to account for these existing tariff preferences when constructing retaliatory tariffs. By incorporating the effects of free trade agreements (FTAs), our dataset measures applied (implemented) tariff rates. For example, Mexico and Canada imposed zero tariffs on U.S. exports in 2017 under the North American Free Trade Agreement. We construct tariff rates as follows. Baseline applied tariff rates at the HS6 level are trade-weighted averages across national tariff lines, taken from TRAINS (following Amiti et al. (2019)). We then add any additional retaliatory increments imposed after 2017. By fixing weights to the pre-war period, our tariff measure captures statutory policy variation without being affected by endogenous shifts in trade composition induced by the tariffs themselves.

The monthly administrative U.S. export records are from the U.S. Census Bureau, which reports export values and quantities at the ten-digit Harmonized System (HS10) level for all trading partners. Our sample covers January 2017 to January 2020 and includes the universe of HS10 codes and destination countries. Because retaliatory tariffs are set at the HS6 level, a concern is that governments may strategically choose which HS6 lines to target based on economically salient trade flows. To mitigate this potential endogeneity, we adapt the exclusion strategy used in Boehm et al. (2023) and remove treated HS10 products that account for all of their destination–HS6 category's trade value in 2017, removing the narrow

[5] Romalis (2007) also uses implemented tariff rates from TRAINS for their analysis.

product lines most likely to have influenced tariff-setting decisions or to reflect targeted protection of narrowly defined, high-value product lines.[6] We further exclude products from non-retaliating countries that experience variation in tariff rates over the sample period, as these changes could be endogenous to the trade conflict and would contaminate the control group by introducing spurious within-product tariff variation.

The input–output (I–O) linkages at the four-digit North American Industry Classification System (NAICS-4) level are from the 2016 "use" table from the Bureau of Economic Analysis (BEA) national input–output accounts.[7] Using the Leontief inverse of the input share matrix, we compute gross input shares of fossil fuels by sector, which incorporate both direct and indirect energy use. On average across tradable sectors, direct fossil fuel input shares are 1.1 percent, while gross shares are 4.3 percent, implying that I–O linkages roughly quadruple sectors' exposure to fossil fuel price changes.[8] Finally, the fossil fuel price index is a composite of petroleum and coal price indices from the Bureau of Labor Statistics (BLS). All datasets are linked through the concordance between the 2017 Harmonized System (HS) and NAICS codes provided by the Census Bureau.

4.1 Event Studies

We start by presenting event studies to visualize the dynamic effects of retaliatory tariffs on U.S. export values, quantities, and prices. The goal is to compare how treated varieties adjust before and after the imposition of retaliatory tariffs relative to untreated varieties.

[6]We aggregate monthly exports over 2017 to construct pre-war export values for each destination–HS10 pair and compute each product's share within its destination–HS6 group. We exclude treated HS10 products whose 2017 share equals one, which are products that are the sole variety within their destination–HS6 group. This restriction removes approximately 1.4% of observations. It ensures that our estimates are not driven by highly salient products that may have influenced tariff-setting decisions. Table B.8 shows that the results are robust to a broader restriction that excludes the dominant HS10 product within each destination–HS6 group when its share exceeds the 25th percentile of the share distribution (approximately 0.95), applied to treated products.

[7]Fajgelbaum et al. (2020) also use U.S. export data from the Census Bureau and I–O linkages from the BEA.

[8]We exclude one sector (sector 34) whose gross fossil fuel input share exceeds unity, reflecting an unusually high energy intensity that would exert undue leverage on the instrument.

Formally, we estimate the following specification:

$$y_{it} = \delta_i + \delta_t + \sum_{j=-12}^{12} \beta_j \, \mathbf{1}(\text{event}_{it} = j) + \varepsilon_{it}, \tag{5}$$

where y_{it} denotes either the logarithm of the ex-tariff price (i.e., the price excluding tariffs, when $y = p$) or the quantity (when $y = q$) of variety i in month t. A variety is defined as an importer–HS10 product pair (with the exporter fixed as the United States throughout). The indicator variable $\mathbf{1}(\text{event}_{it} = j)$ is equal to one if observation (i, t) occurs j months relative to the first month in which an additional retaliatory tariff is imposed. We include variety fixed effects (δ_i) to absorb time-invariant differences across importer-product pairs and month-year fixed effects (δ_t) to capture global shocks common to all trade flows. Standard errors are clustered by country–HS6.[9]

The coefficients β_j measure deviations in outcomes j months before and after the imposition of retaliatory tariffs, relative to the month immediately preceding treatment ($j = -1$), which serves as the omitted reference period. For varieties that never face an additional retaliatory tariff during our sample, all event-time indicators are set to zero (across all months), so these observations serve as pure controls (never-treated control group). Because both variety and time fixed effects are included, the β_j coefficients are identified from differential within-variety changes over time, drawing on comparisons between treated and never-treated varieties as well as across varieties with different treatment timings. The latter comparisons can generate bias under heterogeneous treatment effects, which motivates the Sun-Abraham Sun and Abraham (2021) estimator reported alongside the standard TWFE results.

Figures 3 and 4 display the dynamic responses of U.S. export quantities and prices around the introduction of retaliatory tariffs. As in Figure III of Fajgelbaum et al. (2020), there is no clear evidence of anticipatory behavior prior to the imposition of tariffs: the estimated

[9] The event study uses month-year rather than importer-month fixed effects to preserve cross-country comparisons in a transparent reduced-form setting; the structural estimates are robust to both specifications (Table B.7).

coefficients for the twelve months preceding the event fluctuate around zero.

Following the imposition of retaliatory tariffs ($j = 0$), export quantities decline sharply and persistently, falling by approximately 15 to 25 percent within six months relative to pre-treatment levels. These magnitudes are similar to those reported by Fajgelbaum et al. (2020), who find that quantities decline on average by 25 percent. The similarity between the standard two-way fixed effects (TWFE) and the Sun-Abraham estimators indicates that treatment effect heterogeneity does not drive the results. The Poisson specification in Figure 4 confirms these findings in a framework that accommodates zero trade flows, yielding somewhat larger declines of 18 to 30 percent.[10]

The average post-event effect on export prices is small and statistically insignificant (Figure 3). Retaliatory tariffs do not lead to measurable reductions in U.S. exporter prices. Our results are consistent with broader evidence. Amiti et al. (2019) (Table 3) find no decline in U.S. export prices in response to foreign tariffs. Similarly, Fajgelbaum et al. (2020) observe no change in the before-duty prices. Finally, Cavallo et al. (2021) find an average pass-through of around one-third; their results are not directly comparable as the analysis relies on BLS export price indices aggregated at a coarser sector level.

Average pre-trends may mask differential trends across products exposed to different added rates. If such differential trends were captured by the tariff polynomial, they could mimic tax-dependent compliance. Following Callaway et al. (2024), we re-estimate the event study separately for products in the bottom and top halves of the treatment intensity distribution among treated products, retaining all untreated varieties as controls in each subsample. Treatment intensity is measured as the ratio of the additional to the total tariff rate. Figure B.1 shows no evidence of differential pre-trends across the two groups in the twelve months prior to treatment: pre-treatment coefficients are close to zero in both subsamples, alleviating the concern that the estimated tariff polynomial captures pre-existing differential trends correlated with tariff intensity. Consistent with differences in treatment intensity, we

[10]Implied percentages calculated using $100 \times (e^{\hat{\beta}} - 1)$ for $\hat{\beta} = -0.2$ and $\hat{\beta} = -0.35$.

observe stronger effects in the subsample with above-median treatment intensity.

5 Empirical Specification

In this section, we outline the empirical framework to obtain a consistent estimator of compliance, $\lambda(\cdot)$. We estimate the structural demand equation directly using an instrumental variables strategy, which allows for direct estimation of compliance as a function of tariff rates.

5.1 General Specification

Consider the following demand equation:

$$q_{it} \;\; = \;\; \varepsilon^d p_{it} + \gamma^d T_{it} + \rho^d Z_{it} + \delta_i + \delta_{ct} + u_{it}^d,$$

where i indexes varieties and t indexes months. As before, a variety is defined as an importer–HS10 product pair, so $i \equiv (c, h)$, with c denoting the importing country and h the HS10 product. The dependent variable q_{it} is the log quantity of U.S. exports of variety i in month t, and p_{it} is the corresponding tariff-exclusive log unit value. The tariff variable is $T_{it} = \ln(1 + \tau_{it})$, where τ_{it} denotes the ad valorem tariff rate faced by variety i in month t (measured at the HS6 level and assigned to HS10 products). The variable Z_{it} is a supply shifter. Our primary specification includes variety fixed effects, δ_i, and importer–month fixed effects, δ_{ct}. In this setup, the tariff-exclusive price p_{it} is endogenous, and Z_{it} serves as its instrument.

Following our conceptual framework in Section 3, we impose two restrictions. First, the tariff-induced demand shift satisfies $\gamma^d = \varepsilon^d \lambda(T_{it})$, where $\lambda(T)$ denotes the compliance function. Second, demand is orthogonal to the supply shifter, so $\rho^d = 0$, reflecting the exclusion restriction (SER2) in Section 3. Substituting these restrictions into the demand equation yields

$$q_{it} = \varepsilon^d p_{it} + \varepsilon^d \lambda(T_{it})\, T_{it} + \delta_i + \delta_{ct} + u^d_{it}. \tag{6}$$

To allow compliance to vary flexibly with tariff levels, we approximate $\lambda(T)$ by an order-K polynomial,

$$\lambda(T) = \sum_{k=0}^{K} \theta_k T^k. \tag{7}$$

5.2 Model Selection and Baseline Specification

To select the polynomial order K for the compliance function $\lambda(T)$, we employ repeated out-of-sample cross-validation. We estimate the model for $K = 0, \ldots, 3$ on a randomly selected training half of the data and evaluate predictive accuracy on the held-out test half, repeating this procedure 100 times with different random splits. Table B.1 reports the mean and standard deviation of the out-of-sample MSE across splits for each polynomial order. A quadratic compliance function ($K = 2$) achieves the lowest mean out-of-sample MSE.[11] Moreover, mean MSE declines monotonically from $K = 0$ to $K = 2$, and then rises at $K = 3$, consistent with mild overfitting from unnecessary additional flexibility. This pattern is stable across all 100 repeated splits. We thus adopt the quadratic form for $\lambda(T)$, which allows the rate of decline in compliance to vary with the tariff level while remaining parsimonious:

$$\lambda(T) = \theta_0 + \theta_1 T + \theta_2 T^2. \tag{8}$$

Substituting (8) into the demand equation yields a cubic specification in T_{it}:

$$q_{it} = \varepsilon^d p_{it} + \varepsilon^d \theta_0 T_{it} + \varepsilon^d \theta_1 T^2_{it} + \varepsilon^d \theta_2 T^3_{it} + \delta_i + \delta_{ct} + u^d_{it}. \tag{9}$$

[11] The out-of-sample MSE is evaluated on Equation (6); the model is estimated by IV on the training half, and predictions in the test half are formed using the IV coefficient estimates applied to the observed regressors, including the absorbed fixed effects. We also conducted cross-validation on a reduced-form analog of Equation (6) in which the instrument Z_{it} replaces p_{it} as a regressor and no instrumentation is performed. The MSE differences across polynomial orders are small relative to the variation across splits in both procedures, confirming that the data cannot sharply distinguish among specifications. We adopt the quadratic on parsimony grounds; see Section 6 for sensitivity to this choice.

Our estimation of Equation (9) proceeds via two-stage least squares, instrumenting p_{it} with the supply shifter Z_{it}. Since higher-order terms in T_{it} are deterministic functions of the tariff rate, they are subject to the same endogeneity concerns as T_{it} itself, which we address through the sample restrictions described in Section 4. This strategy identifies ε^d from the coefficient on p_{it}, and we estimate θ_k using $\hat{\theta}_k = \widehat{\varepsilon^d\theta_k}/\hat{\varepsilon}^d$. Finally, we infer the marginal effect of an increase in the tariff rate τ on the compliance rate $\lambda(T)$ from Equation (8):

$$\frac{d\lambda(T)}{d\tau} = \frac{\theta_1}{1+\tau} + \frac{2\theta_2 \ln(1+\tau)}{1+\tau}. \tag{10}$$

5.3 Identification

For the supply shifter Z_{it}, we adopt a Bartik (shift-share) design that measures variation in production energy costs. The share component is the predetermined sector-level fossil fuel input share in 2016 at the NAICS-4 level, and the shift component is the monthly fossil fuel price index. This construction follows the logic in Kilian (2008) that energy price increases raise marginal costs in proportion to energy cost shares. Because these U.S. energy-cost shocks are plausibly orthogonal to demand conditions in a specific foreign market, conditional on variety and importer (country)–month–year fixed effects, the shifter satisfies the exclusion restriction (SER2) in Section 3. Its role is to generate exogenous variation in equilibrium prices that identifies the demand elasticity. Under full compliance, tax variation alone is sufficient for this purpose, because taxes shift demand by a known magnitude according to the size of the tax (Zoutman et al., 2018). In that case, the wedge between the supply and demand functions is known, and tax variation traces out both functions. With imperfect compliance, however, the demand response to tax changes confounds the demand elasticity with the compliance rate, requiring an independent supply shifter to disentangle the two.[12]

A potential concern is that energy prices may proxy for broader global business-cycle

[12] This estimation strategy remains valid under imperfect competition, as shown in Appendix A.2, because conduct parameters are absorbed into the error terms without affecting the slope coefficients that identify ε^d and $\lambda(T)$.

conditions that also affect foreign demand for U.S. exports. Our importer–month fixed effects absorb destination-specific macro shocks (including exchange-rate movements and aggregate demand conditions) that are common across products within a country-month. The remaining threats to identification would require product- or industry-specific demand shifts that systematically covary with energy prices and with the cross-sectional energy-intensity shares used in Z_{it}.

Several features of our design mitigate this concern. First, the energy-intensity shares are measured in 2016, prior to our sample period, and are therefore predetermined with respect to trade-war dynamics. Second, because the monthly fossil-fuel price index is common across products, identification comes from differential exposure driven by cross-sectional variation in energy intensity; in the shift-share framework, the relevant exclusion restriction concerns whether energy-intensity shares are systematically correlated with product-specific demand shocks, conditional on fixed effects (Borusyak et al., 2022). Third, we verify that our estimates are stable when including industry–month-year fixed effects, which absorb demand shocks common to products within broad industry categories. The stability of the estimates across these specifications suggests that product-specific demand confounds correlated with energy intensity are unlikely to drive our results.

To provide more direct evidence on the validity of the instrument, we also implement a pre-period placebo test in the spirit of Borusyak et al. (2022). Specifically, we regress first-differenced log bilateral export quantities on the fossil fuel gross input shares using data from 2014–2015, well before the onset of the trade war. If the shares were correlated with product-specific demand trends, they would predict pre-period export growth. Panel A of B.2 shows that the estimated coefficient on the fossil input shares is 0.029 with a standard error of 0.081, which is economically small and statistically insignificant (N = 1,957,670). This result confirms that the cross-sectional energy-intensity shares do not predict differential bilateral export growth across products within a destination in the pre-period, supporting the maintained assumption that the shares are orthogonal to product-specific demand shocks conditional on the fixed effects included in our main specification.

To further assess which sectors drive this result, we decompose the pre-period estimate following the logic of Goldsmith-Pinkham et al. (2020). The overall coefficient can be written as $\hat{\beta} = \sum_k w_k \hat{\beta}_k$, where w_k is sector k's share of the total within-country variance of the fossil input shares and $\hat{\beta}_k$ is the sector-specific pre-trend slope. Panel B of Table B.2 reports this decomposition. The top 10 sectors account for 68% of the total weight, with no single sector exceeding 15%. The sector-specific estimates $\hat{\beta}_k$ are imprecisely estimated and dispersed around zero, and no individual sector contributes more than 0.06 to the overall coefficient. The diversity of sectors receiving substantial weight, ranging from industrial chemicals to agricultural products and minerals, together with the absence of any dominant sector, supports the view that identification does not hinge on idiosyncratic conditions in a particular industry.

6 Estimation Results

We begin with baseline IV estimates of the demand elasticity and a tariff-dependent compliance function, which we use to compute pre-war and post-war compliance rates and their marginal responses to tariff changes. We then show robustness to alternative cost shifters and fixed effects, examine heterogeneity by intermediate versus final goods and by product differentiation. Finally, we quantify the implied changes in effective tariffs and tariff revenue gaps.

6.1 Baseline Results

Table 2 reports baseline estimates of the first-stage equation and Equation (8) in Panel A, and estimates of (9) in Panel B. Columns (1) and (2) restrict the sample to varieties exported to retaliating countries; Columns (3) and (4) include all U.S. trading partners. All specifications include country-HS10 product fixed effects and country-month-year fixed effects.

Column (1) reports the first-stage regression. The supply shifter (Share $\times$ $\ln(\text{Price}^f)$ with Price^f denoting the monthly fossil fuel price index) is a strong predictor of prices (F-statistic = 49, well above conventional thresholds). The tariff terms in the first stage are small and, where statistically significant, positive, indicating that tariff increases do not reduce exporter prices. This is consistent with prior evidence that tariff incidence during the trade war fell primarily on importers rather than U.S. exporters (Amiti et al., 2019; Fajgelbaum et al., 2020). Column (2) reports the structural demand equation. The coefficient on log price implies a demand elasticity of -1.7, broadly in line with export demand elasticities of -1 to -1.2 estimated by Amiti et al. (2019) and Fajgelbaum et al. (2020) using similar trade war variation. Dividing the coefficients on the tariff terms by the demand elasticity yields the parameters of the compliance function (θ_0, θ_1, and θ_2 from Equation (8) in Section 5): $\hat{\lambda}(T) = 0.91 - 2.5T + 2.3T^2$. The intercept $\hat{\theta}_0 = 0.91$ implies near-full compliance when tariffs are zero, a natural benchmark providing a validity check on the approach.

The lower panel of Table 2 reports several summary measures of compliance. Our preferred measure is the value-weighted average of product-specific compliance rates, which weights each product by its export value and therefore captures compliance on the average exported dollar.[13] Value-weighted pre-war compliance is $\tilde{\text{Avg}}[\hat{\lambda}_{1i}] = 0.78$ (s.e. = 0.20); post-war compliance falls to $\tilde{\text{Avg}}[\hat{\lambda}_{2i}] = 0.53$ (s.e. = 0.11), a decline of approximately 25 percentage points. Evaluating at the mean tariff rate yields similar magnitudes ($\hat{\lambda}_1 = 0.76$, $\hat{\lambda}_2 = 0.46$). These estimates imply that a one percentage-point tariff increase reduces compliance by 1.3 to 2.1 percentage points, depending on the initial tariff level.[14] Expressed as a percent change in non-compliance, which is the metric used by Fisman and Wei (2004),

[13]Because the compliance function is nonlinear, the average of product-specific compliance rates $E[\lambda(T_i)]$ differs from compliance evaluated at the mean tariff rate $\lambda(\bar{T})$, though the two are similar in practice.

[14]Marginal effects are computed from equation (10) as value-weighted averages across treated observations, using export values as weights. The estimate of -2.1 percentage points corresponds to the pre-war period, when the value-weighted mean tariff rate among treated products was 6.7 percent; the estimate of -1.3 percentage points corresponds to the post-war period, when it was 23.7 percent.

this corresponds to a 3 to 9 percent increase.[15] This range includes the 3 percent estimate in Fisman and Wei (2004), who find that a one percentage-point increase in China's tariff rate is associated with a 3 percent increase in evasion of imports from Hong Kong. Importantly, our estimates capture both illegal evasion and legal avoidance, whereas the gap-based approach primarily captures evasion. Results are similar when including all U.S. trading partners (Columns 3–4), with value-weighted compliance rates of 0.79 (s.e. = 0.19) pre-war and 0.53 (s.e. = 0.11) post-war.

Figure 5 plots the estimated compliance function over the support of observed tariff rates among treated varieties. Compliance declines monotonically from approximately 0.91 at the 5th percentile of the tariff distribution to 0.37 at the 95th percentile, with confidence intervals that exclude full compliance at higher tariff levels.[16]

The estimated pre-war compliance rate of 78 percent implies that roughly one-fifth of statutory tariff liability was not reflected in equilibrium outcomes (though the confidence interval at the pre-war mean is wide and does not exclude full compliance). This pre-war magnitude may appear high for trade among developed economies. However, our estimate captures both illegal evasion and legal avoidance, including duty-drawback provisions, bonded warehouses, free trade zones, and firm-specific exemptions. The heterogeneity results below support this interpretation: compliance is systematically lower for intermediate goods, which are more likely to qualify for duty-free treatment.

Two findings emerge. First, compliance declined substantially during the trade war (from

[15]The percent change in evasion is computed as the ratio of the marginal effect on compliance to the evasion rate. At the pre-war tariff level, $(-2.1)/0.22 \approx 9$ percent; at the post-war level, $(-1.3)/0.47 \approx 3$ percent.

[16]Appendix Tables B.3 and B.4 report estimates under the linear ($K = 1$) and cubic ($K = 3$) compliance functions. Post-war compliance is broadly stable across specifications, whereas pre-war estimates are more sensitive to the polynomial order, reflecting the narrower range of pre-war tariff rates, leaving less variation to pin down the shape of the compliance function at low tariff levels. The linear specification cannot capture nonlinear compliance responses and produces an intercept well below one and a modest compliance decline; the cubic specification produces an intercept well above one, indicating overfitting. These patterns confirm the quadratic as a parsimonious choice that yields an economically intuitive intercept near unity, as confirmed in Table B.5, which re-estimates the baseline quadratic specification imposing the boundary condition $\lambda(0) = 1$. The constrained estimates yield slightly higher compliance levels relative to the baseline, but a similar decline from pre- to post-war.

78 to 53 percent), consistent with the well-documented positive relationship between tariff rates and non-compliance (Fisman and Wei, 2004). Second, our framework recovers not only compliance levels but also marginal effects, enabling direct comparison with prior estimates of how non-compliance responds to tariff changes.

6.2 Robustness

Table B.6 reports estimates using alternative cost shifters. Columns (1)–(4) replace the lagged fossil fuel price with the contemporaneous price; Columns (5)–(8) normalize the price relative to 2016. In both cases, demand elasticities range from approximately -1.5 to -1.7, consistent with the baseline. Value-weighted pre-war compliance ranges from 0.77 (s.e. = 0.20) to 0.93 (s.e. = 0.22), and post-war compliance ranges from 0.52 (s.e. = 0.11) to 0.62 (s.e. = 0.12), all broadly consistent with the baseline estimates of 0.78 and 0.53. The similarity of results across specifications indicates that measurement choices in the instrument construction do not drive our findings.

Table B.7 varies the fixed effects structure. Columns (1)–(4) replace country-month-year fixed effects with month-year fixed effects. Columns (5)–(8) add industry-month-year fixed effects to absorb demand shocks common to products within 15 broad industry categories, directly addressing the identification concern discussed in Section 5. Demand elasticities range from -1.63 to -1.76, consistent with the baseline. Value-weighted pre-war compliance ranges from 0.76 (s.e. = 0.20) to 0.81 (s.e. = 0.21), and post-war compliance ranges from 0.51 (s.e. = 0.12) to 0.54 (s.e. = 0.14), all broadly consistent with the baseline estimates of 0.78 and 0.53. Again, the stability of estimates across both specifications supports the baseline findings.

Table B.8 reports estimates using a broader exclusion restriction, dropping all treated HS-10 products whose 2017 export share within their destination–HS-6 group exceeds the 25th percentile of the share distribution, compared to the baseline, which drops only products that made up the entire HS6 value in 2017. The qualitative pattern remains unchanged: compliance declines substantially from the pre-war to the post-war period, with value-weighted pre-war compliance of 0.84 (s.e. = 0.21) and post-war compliance of 0.57 (s.e. = 0.12), broadly consistent with the baseline estimates of 0.78 and 0.53.

Table B.9 examines heterogeneity across country groups with different institutional environments. We estimate the model separately for the EU and Canada, and for China, Mexico, Russia, and Turkey. The two groups differ in measures of institutional quality. In 2017, the year before the onset of the trade war, the mean Transparency International Corruption Perceptions Index (CPI) score was 65.2 for the EU and Canada group, compared with 34.8 for the second group; moreover, every country in the EU and Canada group scored higher than every country in the comparison group.[17] Pre-war tariffs on treated products average 4.4% for EU and Canada (low enough that pre-war compliance approximates θ_0); tariffs are slightly higher (around 8.0%) for China, Mexico, Russia, and Turkey. Pre-war compliance is higher in the low-corruption group: 1.09 (s.e. = 0.29) for the EU and Canada, compared to 0.74 (s.e. = 0.25) for China, Mexico, Russia, and Turkey. This pattern is consistent with weaker institutional environments allowing greater scope for both illegal evasion and legal avoidance. Compliance declines during the trade war in both groups – when tariffs rise to an average of 29.4% for the EU and Canada and 25.0% for the second group – to 0.87 (s.e. = 0.19) and 0.54 (s.e. = 0.15), respectively.[18] Notably, the magnitude of the decline is similar across groups despite their institutional differences. This pattern supports our interpretation of λ as capturing both illegal evasion and legal avoidance: the level differences across groups

[17]Mexico and Turkey score lower than China on the CPI over this period, despite Mexico's OECD membership and Turkey's customs union with the EU; formal institutional integration may not necessarily translate into customs compliance on the ground Dutt and Traca (2010).

[18]The estimated compliance function for the EU and Canada subgroup has a minimum within the observed tariff range (at approximately 19%), implying that compliance increases with the tariff rate beyond that point, and an intercept slightly above one. Both features reflect imprecise estimation on a small sample, as confirmed by the large standard errors.

are consistent with institutional quality limiting evasion in the steady state.[19]

Spillovers and Misclassification One potential limitation is that importers may avoid tariffs by misclassifying products into untreated HS categories. Since tariffs are set at the HS6 level, misclassification would require finding an untreated HS6 category similar enough to plausibly substitute, a considerably higher bar than shifting within a narrow product group, and one that is unlikely to be widely available. If this occurs, our methodology captures non-compliance within HS6 categories but misses shifting across categories.

In Table B.10, we regress log quantities for untreated HS-10 products on three measures of tariff exposure within the same HS-4 category: (i) an indicator for any product in the same HS4 with retaliatory tariffs imposed, (ii) the natural log one plus the maximum additional rate in the same HS4, and (iii) the natural log of one plust the mean added rate in the HS4.[20] The coefficients are positive but statistically insignificant across all three exposure measures and both samples, providing no evidence of systematic misclassification of treated products into nearby untreated codes.

To assess whether misclassification biases our estimates, Table B.11 re-estimates the baseline specification while controlling for the mean tariff rate in the surrounding HS-4 category (excluding the product's own HS-6). If misclassification were substantially biasing our results, this control should partially offset the bias and alter the compliance estimates. Instead, the coefficient on the HS-4 mean rate is small and statistically insignificant. Value-weighted pre-war compliance is 0.69 (s.e. = 0.18), somewhat below the baseline estimate of 0.78, while post-war compliance is 0.55 (s.e. = 0.12), closely in line with the baseline of 0.53. The implied decline of 14 percentage points is smaller than the baseline decline

[19] Estimated changes in compliance include potential concurrent changes in customs enforcement stringency. If retaliating governments adjusted enforcement in response to tariff rate increases, the decline in compliance may be mitigated in countries where institutional quality leaves more scope for such adjustments. The similarity in the magnitude of the change in compliance across groups suggests that differential concurrent changes in enforcement are unlikely to play a primary role in determining the change in compliance, which is consistent with the relatively short time period considered after the tariff hikes. We cannot rule out that illegal evasion and legal avoidance expanded in parallel across institutional environments.

[20] All measures refer to products in the same HS4, excluding the HS6 of the focal product.

of 25 percentage points. However, the estimated compliance rates in Table 2 are all well within the confidence intervals of the analogous estimates presented in Table B.11, which account for misclassification. Since the misclassification control itself is insignificant, this attenuation likely reflects estimation noise from the more demanding specification rather than true omitted-variable bias. The overall pattern of a substantial decline in compliance during the trade war is thus preserved.

6.3 Intermediate and Final Goods

Intermediate goods are often eligible for duty-free importation through mechanisms such as processing trade, bonded warehouses, and free trade zones. To examine whether compliance differs across product types, we estimate separate demand elasticities and compliance functions for intermediate and final goods using the Broad Economic Categories (BEC) classification.

Table 3 reports the reduced-form estimates. The first stage remains strong (F-statistic = 68). Demand is more elastic for final goods ($\varepsilon^d = -3.1$) than for intermediate goods ($\varepsilon^d = -1.7$), consistent with final goods facing greater substitutability at the consumer level. Table 4 reports the structural parameters. The intercept of the compliance function, $\hat{\theta}_0$, is lower for intermediate goods (0.80) than for final goods (0.98), consistent with intermediate goods being more likely to qualify for duty-free treatment (Grant, 2020). Value-weighted pre-war compliance is 0.90 (s.e. = 0.50) for final goods and 0.67 (s.e. = 0.20) for intermediate goods; post-war compliance falls to 0.52 (s.e. = 0.26) and 0.43 (s.e. = 0.11), respectively. Standard errors are smaller for intermediate goods, reflecting that approximately 75 percent of observations fall in this category. The decline is substantial for both product types, and marginal effects are similar: a one percentage point increase in the tariff reduces compliance by 1.2 to 3.0 percentage points, depending on product type and period.

6.4 Product Differentiation

Differentiated products may be more susceptible to tariff non-compliance because their quantities and values are more difficult to verify at customs (Javorcik and Narciso, 2008). To examine this hypothesis, we estimate separate compliance functions for differentiated and non-differentiated products using the Rauch (1999) classification.[21]

Table 5 reports the reduced-form estimates. The first stage remains strong (F-statistic = 70). Demand elasticities are similar across product types: -1.5 for non-differentiated and -1.6 for differentiated products. Table 6 reports the structural parameter estimates. The intercept $\hat{\theta}_0$ is 0.96 for non-differentiated products and 0.79 for differentiated products, indicating that compliance at low tariff rates is substantially higher for homogeneous goods, consistent with the hypothesis that valuation difficulty facilitates non-compliance. Value-weighted pre-war compliance is 0.95 (s.e. = 0.33) for non-differentiated products and 0.62 (s.e. = 0.27) for differentiated products. The near-unity estimate for non-differentiated goods suggests full compliance when tariffs are low and valuation is straightforward. Post-war compliance falls to 0.84 (s.e. = 0.16) for non-differentiated products and 0.43 (s.e. = 0.17) for differentiated products. Evaluating at the mean tariff rate yields qualitatively similar results. Marginal effects are larger for differentiated products: a one percentage point tariff increase reduces compliance by 1.3 to 1.8 percentage points for differentiated goods, compared to 0.4 to 0.6 percentage points for non-differentiated goods, though the latter are imprecisely estimated.

6.5 Effective Tariffs and Revenue Gaps

Table 7 translates the compliance estimates into effective tariff rates and tariff revenue gaps. Panel (a) reports 2017 (pre-war) figures; Panel (b) reports 2019 (post-war) figures. The table presents two classification schemes: intermediate versus final goods, and differentiated versus non-differentiated products.

[21] Products that cannot be classified under this system are excluded from the analysis.

In 2017, statutory tariffs average 0.9 to 1.5 percent across product categories, but effective tariffs, defined as the product of statutory rates and compliance rates, are substantially lower for categories with compliance below one. The large disparity in export values between intermediate and final goods reflects the composition of U.S. trade, with exports heavily concentrated in upstream sectors such as capital goods, industrial inputs, chemicals, and agricultural commodities rather than finished consumer products. As a result, intermediate goods account for the bulk of both tariff exposure and the revenue gaps reported below. The gap between statutory and effective tariffs implies foregone tariff revenue of $0.2 billion for final goods and $3.7 billion for intermediate goods. For intermediate goods, part of this gap reflects legal duty-free mechanisms rather than enforcement failures. Under the differentiation classification, the revenue gap is $2.7 billion for differentiated products; non-differentiated products exhibit near-full compliance, with a revenue gap of $0.2 billion.[22]

By 2019, statutory tariffs had roughly doubled due to retaliatory measures, but compliance had fallen sharply. Revenue gaps rise to $0.6 billion for final goods and $9.0 billion for intermediate goods. Under the differentiation classification, revenue gaps are $4.8 billion for differentiated products and $0.9 billion for non-differentiated products. Column (6) isolates the portion attributable to the trade war, that is, the additional revenue gap caused by declining compliance rates, holding 2019 tariffs and trade values fixed. Under the intermediate/final classification, war-induced revenue gaps total $0.3 billion (final) and $3.2 billion (intermediate), or $3.5 billion combined, roughly 36 percent of the total revenue gap in 2019. Under the differentiation classification, war-induced gaps total $1.3 billion (differentiated) and $0.5 billion (non-differentiated), or $1.8 billion combined, approximately 32 percent of the total revenue gap.[23]

These magnitudes underscore that higher statutory rates are associated with greater

[22] The revenue loss (Column 5) is smaller under the differentiation classification ($2.9 billion versus $3.9 billion) because products that cannot be classified under Rauch (1999) are excluded from that sample.

[23] The war-induced revenue gap is smaller under the differentiation classification ($1.8 billion versus $3.5 billion) primarily because products that cannot be classified under Rauch (1999) are excluded from that sample, reducing coverage to approximately 70 percent of total export values.

departures from full implementation, attenuating the revenue and protective effects of retaliatory tariffs. Because our compliance measure captures both illegal evasion and legal avoidance mechanisms that governments have deliberately designed to reduce effective tariff burdens on export-oriented firms, the revenue gaps in Table 7 should be interpreted as reflecting a combination of enforcement shortfalls and intended policy design rather than enforcement failure alone.

7 Conclusion

This paper develops a novel method to estimate tax compliance rates from variation in statutory tax rates using market data on prices and quantities. The key insight is that compliance can be separately identified from demand elasticities by comparing quantity responses to tax changes with responses to an independent supply shifter. Our approach allows compliance to vary with the tax rate, accommodating the well-documented positive relationship between tax rates and non-compliance.

We apply the method to retaliatory tariffs imposed on U.S. exports during the 2018 trade war. Compliance fell from 78 percent at pre-war tariff rates (averaging 6 percent among treated products) to 53 percent at post-war rates (averaging around 20 percent). The implied marginal effects are comparable in magnitude to estimates from the gap-based approach of Fisman and Wei (2004). Importantly, our estimates capture both illegal evasion and legal avoidance mechanisms such as duty-free zones, bonded warehouses, and processing trade regimes. Heterogeneity analysis supports this interpretation: compliance is systematically lower for intermediate goods, which are more likely to qualify for duty-free treatment, and for differentiated products, whose values are harder to verify at customs.

The results have implications for tariff policy. We estimate that the decline in compliance during the trade war reduced tariff revenue for U.S. trading partners by approximately $3.5 billion relative to pre-war compliance rates. This suggests that the effective incidence of tariff increases was substantially smaller than statutory rates imply, attenuating both the protective effect and the revenue yield of trade war tariffs.

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Tables and Figures

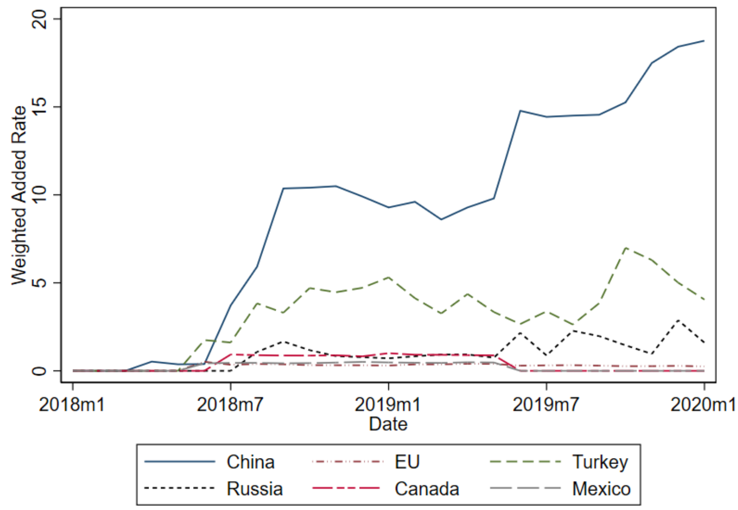
Table 1: The 2018-2019 Trade War

Country	Month enacted	Products	2017 exports		Tariff(%)		
		(#HS-10)	(million $US)	(%)	2017	2018	2019
China	April, 2018	6,241	87,522	8.1	7.5	11.9	23.8
Mexico	June, 2018	115	4,500	0.4	0	12.6	9.2
Turkey	June, 2018	158	1,493	0.1	11.5	19.2	45.8
European Union	June, 2018	188	3,950	0.4	3.1	14.5	27.6
Canada	July, 2018	356	18,793	1.7	0	6.1	6.0
Russia	August, 2018	117	253	0.0	2.9	11.7	33.0
Total		7,175	116,511	10.7	5.9	11.1	20.4

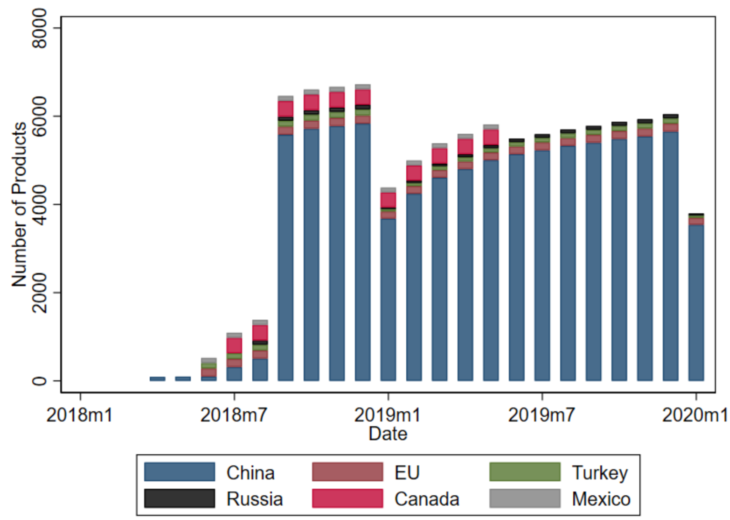
Notes: The table summarizes the timeline, scope, and tariff rates of retaliatory measures imposed on U.S. exports in 2018–2019. Products refers to the number of importer-HS-10 product pairs targeted. Export values and shares are based on 2017 U.S. export data; the share denominator is total 2017 U.S. exports. Tariff rates are annual averages across retaliating HS-10 products weighted by 2017 trade values. Data on retaliatory tariff schedules are from Cavallo et al. (2021); baseline applied tariffs are from TRAINS; U.S. export values are from the U.S. Census Bureau.

Figure 1: Retaliating Tariffs and Coverage

(a) Weighted Average Added Rates

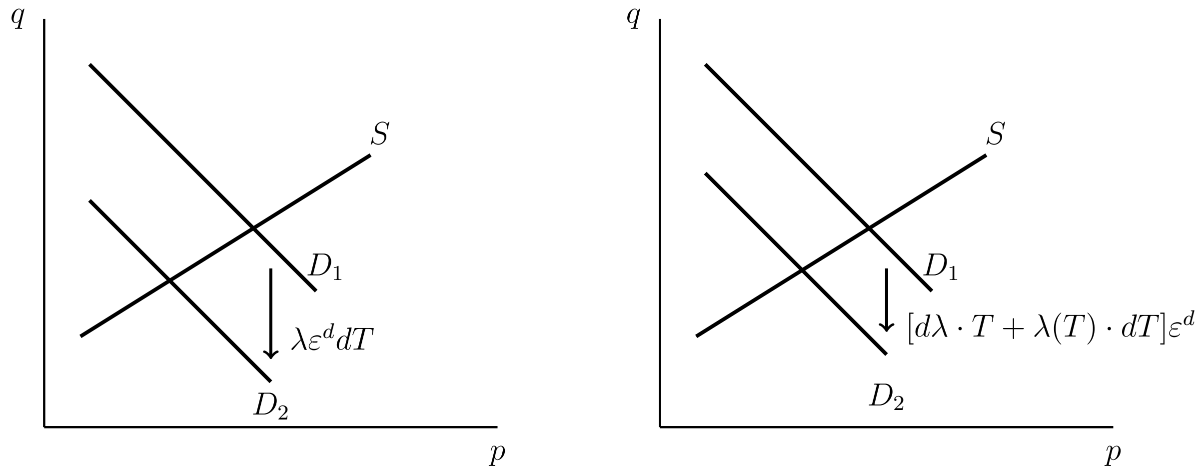


(b) Number of Products in Retaliation



Notes: Panel (a) shows the average added tariff rates weighted by trade values for each retaliating country, from January 2018 to January 2020. Countries are ordered by added tariffs in January 2020 from high to low: China, Turkey, Russia, EU, Canada, and Mexico. Panel (b) shows the number of HS-10-level products covered under retaliatory tariffs. Each importer-HS-10 pair is counted as a product variety.

Figure 2: Estimating Compliance with a Tax Increase

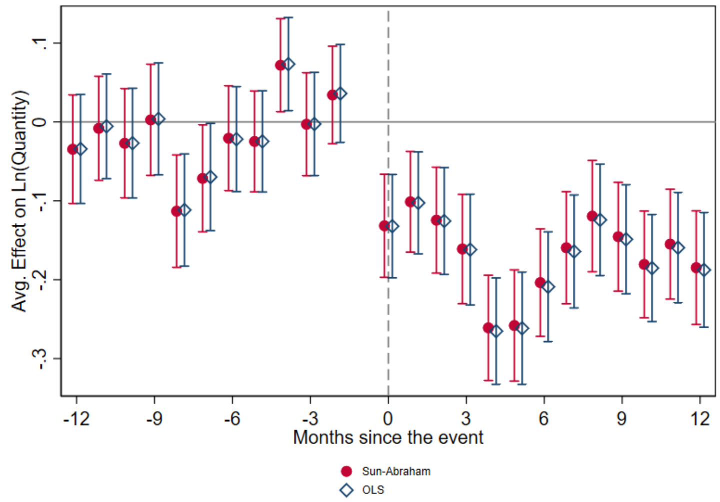


(a) Compliance is Constant in the Tax

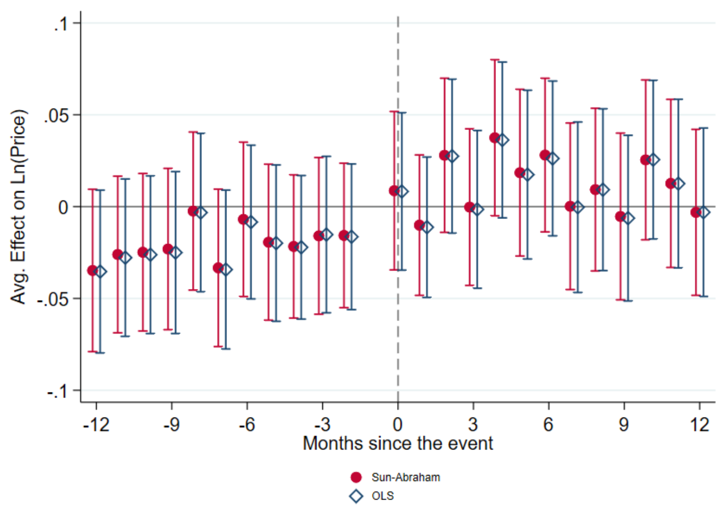
(b) Compliance Depends on the Tax

Notes: The figure illustrates how demand shifts in response to a tariff increase when the statutory burden falls on the demand side. Prices are tax-exclusive; demand shifts reflect changes in the tax-inclusive price faced by importers. Panel (a) shows the case of constant compliance: the demand curve shifts down by $\lambda\varepsilon^d dT$, attenuating the shift relative to full compliance by the factor λ. Panel (b) shows the general case in which compliance depends on the tariff level: the demand shift includes both the direct effect of the tariff increase and the induced change in compliance, $[d\lambda \cdot T + \lambda(T) \cdot dT]\varepsilon^d$. When compliance is decreasing in the tariff rate ($d\lambda < 0$), the demand shift is smaller than in Panel (a), implying that higher tariffs are partially self-defeating.

Figure 3: Event Studies, Binary Treatment, Quantity and Price



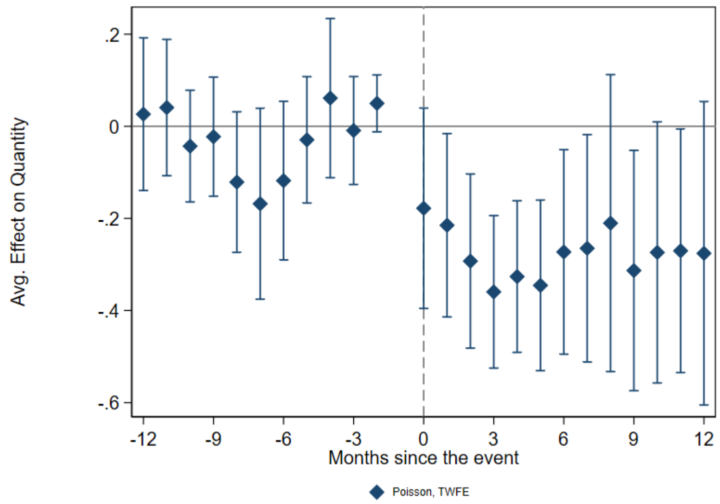
(a) Effects of Added Tariff on ln(Quantity).



(b) Effect of Added Tariff on ln(Price)

Notes: The figure reports event study estimates of the effect of being subject to any retaliatory tariff on log quantity (Panel a) and log price (Panel b). Observations are at the country-HS-10-month level. Each panel reports two sets of estimates: OLS with two-way fixed effects and the heterogeneity-robust estimator of Sun and Abraham (2021). Both include variety (country-HS-10) and month-year fixed effects. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Confidence intervals are based on standard errors clustered at the country-HS-6 level.

Figure 4: Poisson Event Studies, Binary Treatment, Quantity



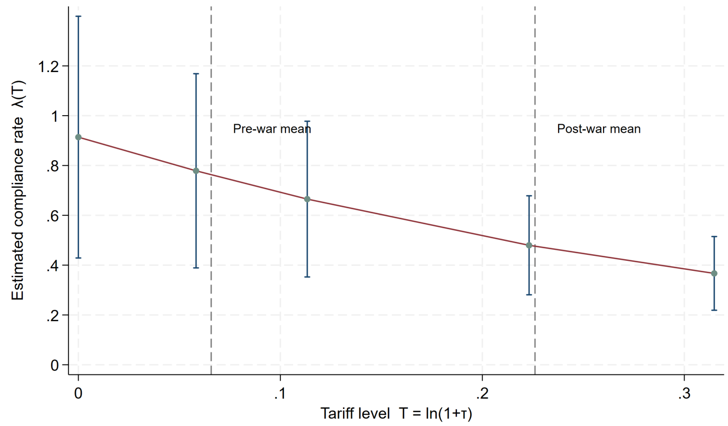
Notes: The figure reports Poisson event study estimates of the effect of being subject to any retaliatory tariff on export quantities. Observations are at the country-HS-10-month level. The specification includes variety (country-HS-10) and month-year fixed effects. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Confidence intervals are based on standard errors clustered at the country-HS-6 level.

Table 2: Tax-Dependent Compliance Rate Estimates

	(1)	(2)	(3)	(4)
	Retaliating Countries Only		*All Sample Countries*	
	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)
T	0.391**	−1.563***	0.396**	−1.569***
	(0.197)	(0.353)	(0.197)	(0.349)
T^2	−1.089	4.195**	-1.105	4.212**
	(0.831)	(1.506)	(0.831)	(1.495)
T^3	0.443	−3.884**	0.455	−3.891**
	(0.856)	(1.580)	(0.857)	(1.573)
Share × ln(Price^f)	4.311***		3.479***	
	(0.617)		(0.419)	
ln($Price$)		−1.711***		−1.696***
		(0.183)		(0.151)
Mean	4.145	7.172	4.109	6.762
F-Stat	48.756		68.957	
Observations	1,488,987	1,488,987	3,399,661	3,399,661
Clusters	54,241	54,241	158,672	158,672
$\hat{\theta}_0$		0.913		0.925
SE($\hat{\theta}_0$)		0.247		0.235
$\hat{\theta}_1$		−2.452		−2.482
SE($\hat{\theta}_1$)		0.958		0.936
$\hat{\theta}_2$		2.27		2.293
SE($\hat{\theta}_2$)		0.972		0.962
$\hat{\lambda}_1$		0.756		0.765
SE($\hat{\lambda}_1$)		0.19		0.18
$\hat{\lambda}_2$		0.464		0.469
SE($\hat{\lambda}_2$)		0.097		0.09
Avg[$\hat{\lambda}_{1i}$]		0.763		0.773
SE(Avg[$\hat{\lambda}_{1i}$])		0.193		0.182
Avg[$\hat{\lambda}_{2i}$]		0.482		0.488
SE(Avg[$\hat{\lambda}_{2i}$])		0.1		0.093
$\tilde{\text{Avg}}$[$\hat{\lambda}_{1i}$]		0.783		0.793
SE($\tilde{\text{Avg}}$[$\hat{\lambda}_{1i}$])		0.2		0.189
$\tilde{\text{Avg}}$[$\hat{\lambda}_{2i}$]		0.527		0.534
SE($\tilde{\text{Avg}}$[$\hat{\lambda}_{2i}$])		0.113		0.106
Country-HS-10 FE	x	x	x	x
Country-Month-Year FE	x	x	x	x

Notes: The table reports first-stage, reduced-form, and structural estimates of the tax-dependent compliance function. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped to address potential endogenous tariff targeting. Post-war periods are pooled. Columns (1) and (2) include varieties exported to retaliating countries only; Columns (3) and (4) include all U.S. trading partners. The structural parameters are recovered as described in Section 5. $\hat{\lambda}$ refers to compliance at the mean tax rate among treated observations. $Avg[\hat{\lambda}_i]$ refers to simple average of product-specific compliance rate estimates. $\tilde{Avg}[\hat{\lambda}_i]$ refers to average of product-specific compliance rate estimates weighted by product value. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Figure 5: Estimated Compliance Function



Notes: The figure plots the estimated compliance function $\hat{\lambda}(T)$, where $T = \ln(1+\tau)$ and τ is the ad valorem tariff rate, evaluated at the 5th, 25th, 50th, 75th, and 95th percentiles of the tariff distribution among treated varieties over the full sample period. Vertical bars denote 95 percent confidence intervals. Dashed vertical lines indicate the pre-war mean tariff ($\bar{T}_{\text{pre}} \approx 0.066$, corresponding to $\bar{\tau}_{\text{pre}} \approx 6.8\%$) and the post-war mean tariff ($\bar{T}_{\text{post}} \approx 0.226$, corresponding to $\bar{\tau}_{\text{post}} \approx 25.4\%$), computed as unweighted means of T across treated variety–month observations before and after the imposition of retaliatory tariffs, respectively. The post-war figure exceeds the value-weighted average of 20.4% reported in Table 1 because unweighted means assign greater weight to high-tariff countries relative to the trade-value-weighted averages. Estimates are based on Column (2) of Table 2.

Table 3: Reduced Form, Intermediate and Final Goods

	(1) *Pooled First Stage* ln(Price)	(2) IV ln(Quantity)
1[Final Good]		
× T	0.314	−3.050**
	(0.481)	(1.489)
× T^2	−1.245	10.275
	(2.258)	(7.155)
× T^3	0.491	−11.271
	(3.117)	(10.563)
1[Intermediate Good]		
× T	0.300	−1.352***
	(0.222)	(0.406)
× T^2	−0.413	3.977**
	(0.964)	(1.851)
× T^3	−0.199	−3.794**
	(0.962)	(1.878)
Share × ln(Price^f)	3.462***	
	(0.419)	
ln(Price)		
× 1[Final Good]		−3.115***
		(0.661)
× 1[Intermediate Good]		−1.700***
		(0.171)
Mean	4.108	6.763
F-Stat	68.186	
Observations	3,398,041	3,398,041
Clusters	158,633	158,633
Country-HS-10 FE	x	x
Country-Month-Year FE	x	x

Notes: The table reports first-stage and reduced-form estimates separately for intermediate and final goods, classified using the Broad Economic Categories (BEC) system. Intermediate goods are more likely to qualify for duty-free treatment through mechanisms such as processing trade, bonded warehouses, and free trade zones; final goods are less likely to qualify. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Column (1) reports the first-stage estimates using a single common shifter. Column (2) reports the second-stage estimates allowing for different demand elasticities across product types, instrumenting ln(price) interacted with product-type indicators using the cost shifter interacted with the same indicators. Structural parameter estimates are reported in Table 4. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 4: Compliance, Intermediate and Final Goods

	(1)	(2)
	$1[FinalGood]$	$1[IntGood]$
$\hat{\theta}_0$	0.979	0.795
$SE(\hat{\theta}_0)$	0.559	0.261
$\hat{\theta}_1$	-3.298	-2.339
$SE(\hat{\theta}_1)$	2.473	1.122
$\hat{\theta}_2$	3.618	2.231
$SE(\hat{\theta}_2)$	3.5	1.123
$\hat{\lambda}_1$	0.709	0.657
$SE(\hat{\lambda}_1)$	0.377	0.201
$\hat{\lambda}_2$	0.34	0.389
$SE(\hat{\lambda}_2)$	0.162	0.1
$\text{Avg}[\hat{\lambda}_{1i}]$	0.726	0.663
$SE(Avg[\hat{\lambda}_{1i}])$	0.389	0.203
$Avg[\hat{\lambda}_{2i}]$	0.366	0.406
$SE(Avg[\hat{\lambda}_{2i}])$	0.174	0.104
$\tilde{\text{Avg}}[\hat{\lambda}_{1i}]$	0.899	0.666
$SE(\tilde{Avg}[\hat{\lambda}_{1i}])$	0.504	0.204
$\tilde{Avg}[\hat{\lambda}_{2i}]$	0.521	0.428
$SE(\tilde{\text{Avg}}[\hat{\lambda}_{2i}])$	0.262	0.11

Notes: The table reports structural estimates of the compliance function separately for final and intermediate goods, based on the IV estimates in Column (2) of Table 3. Intermediate goods are more likely to qualify for duty-free treatment through mechanisms such as processing trade, bonded warehouses, and free trade zones; final goods are less likely to qualify. Classifications follow the Broad Economic Categories (BEC) system. Structural parameters are recovered as described in Section 5. $\hat{\lambda}$ refers to compliance at the mean tax rate among treated observations. $Avg[\hat{\lambda}_i]$ refers to simple average of product-specific compliance rate estimates. $\tilde{Avg}[\hat{\lambda}_i]$ refers to average of product-specific compliance rate estimates weighted by product value. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors are clustered at the country-HS-6 level.

Table 5: Reduced Form, By Product Differentiation

	(1) *Pooled First Stage*	(2) IV, Different ε_d
	ln(Price)	ln(Quantity)
1[Non-Differentiated]		
$\times\ T$	0.164	-1.440^{**}
	(0.221)	(0.500)
$\times\ T^2$	-0.785	0.973
	(0.919)	(2.084)
$\times\ T^3$	0.189	-0.664
	(0.901)	(2.002)
1[Differentiated]		
$\times\ T$	0.635^{**}	-1.240^{**}
	(0.317)	(0.488)
$\times\ T^2$	-1.932	3.525^{*}
	(1.316)	(2.000)
$\times\ T^3$	1.264	-2.721
	(1.319)	(2.038)
Share $\times$ ln(Price^f)	3.505^{***}	
	(0.419)	
ln(Price)		
$\times$ 1[Non-Differentiated]		-1.494^{***}
		(0.168)
$\times$ 1[Differentiated]		-1.574^{***}
		(0.211)
Mean	4.109	6.762
F-Stat	69.929	
Observations	3,399,661	3,399,661
Clusters	158,672	158,672
Country-HS-10 FE	x	x
Country-Month-Year FE	x	x

Notes: The table reports first-stage and reduced-form estimates separately for non-differentiated and differentiated products, classified using Rauch (1999). Differentiated products lack observable reference prices, making customs valuation harder and creating greater scope for non-compliance. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Column (1) reports the first-stage estimates using a single common shifter. Column (2) reports the second-stage estimates allowing for different demand elasticities across product types, instrumenting ln(price) interacted with product-type indicators using the cost shifter interacted with the same indicators. Structural parameter estimates are reported in Table 6. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6: Compliance, By Product Differentiation

	(1) Non-Differentiated	(2) Differentiated
$\hat{\theta}_0$	0.963	0.787
$SE(\hat{\theta}_0)$	0.359	0.358
$\hat{\theta}_1$	-0.651	-2.239
$SE(\hat{\theta}_1)$	1.401	1.37
$\hat{\theta}_2$	0.444	1.728
$SE(\hat{\theta}_2)$	1.341	1.347
$\hat{\lambda}_1$	0.927	0.64
$SE(\hat{\lambda}_1)$	0.292	0.275
$\hat{\lambda}_2$	0.832	0.363
$SE(\hat{\lambda}_2)$	0.155	0.138
$Avg[\hat{\lambda}_{1i}]$	0.929	0.646
$SE(Avg[\hat{\lambda}_{1i}])$	0.296	0.279
$Avg[\hat{\lambda}_{2i}]$	0.837	0.376
$SE(Avg[\hat{\lambda}_{2i}])$	0.159	0.143
$\tilde{Avg}[\hat{\lambda}_{1i}]$	0.945	0.621
$SE(\tilde{Avg}[\hat{\lambda}_{1i}])$	0.326	0.265
$\tilde{Avg}[\hat{\lambda}_{2i}]$	0.841	0.432
$SE(\tilde{Avg}[\hat{\lambda}_{2i}])$	0.163	0.168

Notes: The table reports structural estimates of the compliance function for non-differentiated and differentiated products, based on the IV estimates in Column (2) of Table 5. Product classification follows Rauch (1999). We use the instrument Z interacted with product-type indicators to instrument for ln(price) interacted with the same indicators. Structural parameters are recovered as described in Section 5. $\hat{\lambda}$ refers to compliance at the mean tax rate among treated observations. $Avg[\hat{\lambda}_i]$ refers to simple average of product-specific compliance rate estimates. $\tilde{Avg}[\hat{\lambda}_i]$ refers to average of product-specific compliance rate estimates weighted by product value. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 7: Effective Tariffs by Product Type

	Panel (a): Compliance and Effective Tariffs in 2017					
	(1) Statutory Tariffs (%)	(2) Compliance Rate	(3) Effective Tariffs (%)	(4) U.S. Export Values ($B)	(5) Revenue Loss ($B)	
Final Goods	0.935	0.952	0.691	69.9	0.2	
Intermediate Goods	1.404	0.767	0.809	625.2	3.7	
Differentiated Products	1.515	0.758	0.826	389.2	2.7	
Non-differentiated Products	1.310	0.956	1.178	115.7	0.2	
	Panel (b): Compliance and Effective Tariffs in 2019					
	(1) Statutory Tariffs (%)	(2) Compliance Rate	(3) Effective Tariffs (%)	(4) U.S. Export Values ($B)	(5) Revenue Loss ($B)	(6) War-induced Loss ($B)
Final Goods	1.742	0.937	0.953	73.4	0.6	0.3
Intermediate Goods	2.432	0.753	1.100	672.6	9.0	3.2
Differentiated Products	2.242	0.747	1.062	409.0	4.8	1.3
Non-differentiated Products	4.202	0.944	3.464	120.4	0.9	0.5

Notes: The table translates the structural compliance estimates into effective tariff rates and tariff revenue gaps. Each panel presents results under two separate product classifications: intermediate versus final goods (rows 1–2) and differentiated versus non-differentiated products (rows 3–4), following the Broad Economic Categories (BEC) and Rauch (1999) classifications, respectively. Panel (a) reports pre-war figures using 2017 tariff rates and export values; Panel (b) reports post-war figures using 2019 tariff rates and export values. All figures are export-value-weighted averages across the products in each category. Column (1) reports statutory tariff rates averaged across all U.S. exports in each product category (including the large share of products facing zero retaliatory tariffs). Column (2) reports average compliance rates computed based on the structural estimates in Tables 4 and 6. Column (3) reports trade-value-weighted average effective tariff rates, where each product's effective tariff is computed as the product of its own statutory tariff rate and its estimated compliance rate $\hat{\lambda}(T_i)$; because $\hat{\lambda}(T)$ is nonlinear and decreasing in the tariff level, column (3) differs from the product of columns (1) and (2). Column (4) reports observed U.S. export values. Column (5) reports the total tariff revenue gap, defined as the difference between statutory and effective tariff revenue at observed export values. Column (6), available in Panel (b) only, reports the war-induced revenue gap: the additional revenue loss attributable solely to the *decline* in compliance between 2017 and 2019, holding 2019 tariff rates and export values fixed. Revenue gaps capture both illegal evasion and legal avoidance, including duty-drawback provisions, bonded warehouses, free trade zones, and processing trade regimes. Products that cannot be classified under Rauch (1999) are excluded from rows 3–4, reducing coverage to approximately 70 percent of total export values.

Appendix A: Supplemental Results

A.1 Estimation from Changes in Supply-Side Tax Rates

Constant Compliance Case

The base model considers compliance estimation when the statutory burden falls on demand and their tax rate changes; in that case, a supplier shifter is required. In this appendix, we show that a demand shifter is required when the statutory burden instead lies on the supply side. Start with the same model, but letting p_{it} refer to the tax-inclusive price, we have:

$$\begin{aligned} q_{it} &= \varepsilon^d p_{it} + \gamma^d T_{it} + \rho^d Z_{it} + v_{it}^d, \\ q_{it} &= \varepsilon^s p_{it} + \gamma^s T_{it} + \rho^s Z_{it} + v_{it}^s, \end{aligned}$$

where Z_{it} denotes a demand shifter. Consistent with the literature, we assume that the structural equations are written in logarithms; thus, price coefficients are the structural demand and supply elasticities. Equating demand and supply generates the following reduced-form equations (nothing changes yet since there are general relations before imposing assumptions):

$$\begin{aligned} q_{it} &= \frac{\gamma^d \varepsilon^s - \gamma^s \varepsilon^d}{\varepsilon^s - \varepsilon^d} T_{it} + \frac{\rho^d \varepsilon^s - \rho^s \varepsilon^d}{\varepsilon^s - \varepsilon^d} Z_{it} + \zeta_{it}^y, \\ p_{it} &= \frac{\gamma^d - \gamma^s}{\varepsilon^s - \varepsilon^d} T_{it} + \frac{\rho^d - \rho^s}{\varepsilon^s - \varepsilon^d} Z_{it} + \zeta_{it}^p. \end{aligned}$$

Let π_{Tq}, π_{Tp}, π_{Zq}, and π_{Zp} capture the four reduced form coefficients in the two equations above so that

$$\begin{aligned}
\pi_{Tq} &= \frac{\gamma^d \varepsilon^s - \gamma^s \varepsilon^d}{\varepsilon^s - \varepsilon^d}, \\
\pi_{Tp} &= \frac{\gamma^d - \gamma^s}{\varepsilon^s - \varepsilon^d}, \\
\pi_{Zq} &= \frac{\rho^d \varepsilon^s - \rho^s \varepsilon^d}{\varepsilon^s - \varepsilon^d}, \\
\pi_{Zp} &= \frac{\rho^d - \rho^s}{\varepsilon^s - \varepsilon^d}.
\end{aligned}$$

With a tax on the supply side of the market, i.e., sellers bear the statutory burden of the tax. When the tax rate changes, the corresponding shift in supply is scaled by the *constant* level of compliance (λ). In this case the identification of the structural parameters is based on the following assumptions:

Assumption 4. *Exclusion of Tax from Demand Equation: $\gamma^d = 0$. Tax is imposed on the supply side.*

Assumption 5. *Standard Exclusion Restriction (SER2). The variable Z_{it} is a demand shifter and does not appear in the structural supply equation: $\rho^s = 0$.*

Assumption 6. *Ramsey Exclusion Restriction (RER Supply). Tax changes induce supply shifts that depend on compliance and the supply elasticity: $\gamma^s = -\lambda \varepsilon^s$.*

Under Assumptions 4 and 5, we express the supply and demand elasticities as follows:

$$\varepsilon^s = \frac{\pi_{Zq}}{\pi_{Zp}},$$

$$\varepsilon^d = \frac{\pi_{Tq}}{\pi_{Tp}}.$$

In addition, we also have that:

$$\pi_{Tp}(\varepsilon^s - \varepsilon^d) = -\gamma^s$$

Using Assumption 6, which assumes a constant level of compliance, we can solve for λ.

$$\lambda = \pi_{Tp} - \frac{\pi_{Tq}}{\varepsilon^s}.$$

A.2 Imperfect Competition

Following Dearing (2022), we develop demand and profit-maximizing conditions under imperfect competition that generate structural equations that mirror those for demand and supply above. As above, we start with the case of a constant compliance rate. To do so, start with the following demand function:

$$Q = z^{\rho^d} \cdot (T^\lambda P)^{\varepsilon^d}.$$

This produces the following inverse demand:

$$P(Q) = \frac{\left(z^{-\rho^d} \cdot Q\right)^{\frac{1}{\varepsilon^d}}}{T^\lambda}.$$

Lastly, suppose that a firm's profit is given by:

$$\Pi = P(Q)q - c(Z_{it})q^\phi,$$

with $\phi > 1$ to ensure that the second-order conditions of profit maximization hold under the CES demand, and $c(Z_{it})$ denoting a cost shifter that depends on the supply shifter Z_{it}.

We use the conduct parameter approach, where θ refers to conduct. The first-order condition for profit maximization is:

$$P - c(Z_{it})\phi q^{\phi-1} + \frac{P\theta}{Q\varepsilon^d} q = 0.$$

This nests Bertrand competition ($\theta = 0$), Cournot competition ($\theta = 1$), and collusion ($\theta =$

N). In a symmetric equilibrium with N firms, $q = Q/N$, so the first-order condition becomes:

$$P\left(1+\frac{\theta}{N\varepsilon^d}\right) \;=\; c(Z_{it})\phi\left(\frac{Q}{N}\right)^{\phi-1}.$$

This produces our two structural equations, which are analogous to the demand and supply equations in the base model but derived from profit maximization under imperfect competition:

$$\begin{aligned} Q &= z^{\rho^d}\cdot[T^\lambda P]^{\varepsilon^d}, \\ P\left(1+\frac{\theta}{N\varepsilon^d}\right) &= c(Z_{it})\phi\left(\frac{Q}{N}\right)^{\phi-1}. \end{aligned}$$

Taking logs of the second equation and rearranging:

$$\ln P + \underbrace{\ln\left(1+\frac{\theta}{N\varepsilon^d}\right)}_{\text{constant}} = \underbrace{\ln(c(Z_{it})\phi)}_{\text{constant}-\rho^s Z_{it}} + (\phi-1)\left[\ln Q - \underbrace{\ln N}_{\text{constant}}\right].$$

Rearranging to isolate $\ln Q$, and collecting all terms that do not vary with the regressors into a single constant κ:

$$\begin{aligned} (\phi-1)\ln Q &= \ln P + \rho^s Z_{it} + \underbrace{\ln\left(1+\frac{\theta}{N\varepsilon^d}\right) - \ln(c_0\phi) + (\phi-1)\ln N}_{\equiv\kappa}, \\ \ln Q &= \frac{1}{\phi-1}\ln P + \frac{\rho^s}{\phi-1}Z_{it} + \frac{\kappa}{\phi-1}, \end{aligned}$$

where c_0 is the Z_{it}-invariant component of marginal cost. The Z_{it}-varying component, $\rho^s Z_{it}$, carries through to the panel supply-equivalent equation, providing the supply shifter that enters as $\frac{\rho^s}{\phi-1}Z_{it}$ in the system below.

Moving to panel notation, replacing $\ln Q$ with q_{it} and $\ln P$ with p_{it}, and absorbing $\frac{\kappa}{\phi-1}$ into the variety fixed effects, we obtain the supply-equivalent equation. The demand equation follows directly from taking logs of $Q = z^{\rho^d}(T^\lambda P)^{\varepsilon^d}$. All conduct and market structure terms

(θ, N, and c) are absorbed into the fixed effects, leaving the slope coefficients unaffected by market structure. We therefore obtain the following system that mirrors the structure of the competitive case:

$$
\begin{aligned}
q_{it} &= \varepsilon^d p_{it} + \lambda \varepsilon^d T_{it} + \rho^d Z_{it} + \psi_{it}^d, \\
q_{it} &= \frac{1}{\phi - 1} p_{it} + \frac{\rho^s}{\phi - 1} Z_{it} + \psi_{it}^s.
\end{aligned}
$$

Let $\rho^d = 0$ so that Z_{it} acts as a supply shifter and is omitted from the demand equation. Let $\phi^s = \frac{1}{\phi-1}$. We can then write the reduced form equations as:

$$
\begin{aligned}
p_{it} &= \frac{\lambda \varepsilon^d}{\phi^s - \varepsilon^d} T_{it} - \frac{\phi^s \rho^s}{\phi^s - \varepsilon^d} Z_{it} + \psi_{it}^p, \\
q_{it} &= \frac{\phi^s \lambda \varepsilon^d}{\phi^s - \varepsilon^d} T_{it} - \frac{\phi^s \rho^s \varepsilon^d}{\phi^s - \varepsilon^d} Z_{it} + \psi_{it}^y,
\end{aligned}
$$

The reduced form coefficients are:

$$
\begin{aligned}
\pi_{Tp} &= \frac{\lambda \varepsilon^d}{\phi^s - \varepsilon^d}, \\
\pi_{Zp} &= -\frac{\phi^s \rho^s}{\phi^s - \varepsilon^d}, \\
\pi_{Tq} &= \frac{\phi^s \lambda \varepsilon^d}{\phi^s - \varepsilon^d}, \\
\pi_{Zq} &= -\frac{\phi^s \rho^s \varepsilon^d}{\phi^s - \varepsilon^d}.
\end{aligned}
$$

This produces the following estimators:

$$
\begin{aligned}
\varepsilon^d &= \frac{\pi_{Zq}}{\pi_{Zp}}, \\
\phi^s &= \frac{\pi_{Tq}}{\pi_{Tp}}, \\
\lambda &= -\pi_{Tp} + \frac{\pi_{Tq}}{\varepsilon^d}.
\end{aligned}
$$

A.2.1 Estimation with Tax-Dependent Compliance

As in the base model, we replace λ with $\lambda(T)$ so that

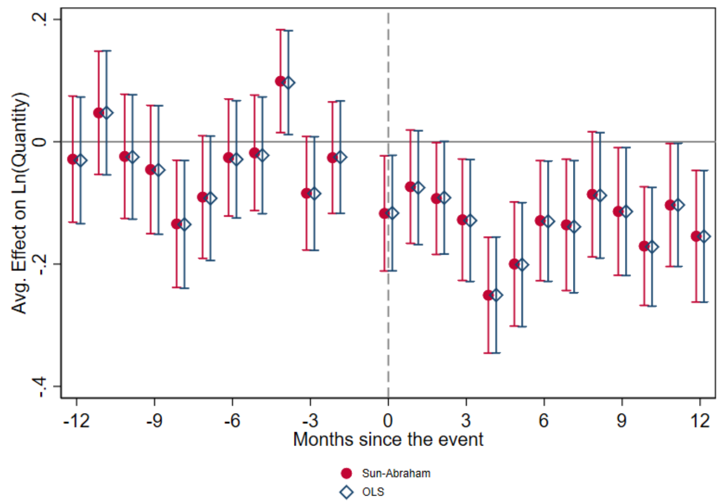
$$\begin{aligned}
p_{it} &= \frac{\lambda(T)\varepsilon^d}{\phi^s - \varepsilon^d} T_{it} - \frac{\phi^s \rho^s}{\phi^s - \varepsilon^d} Z_{it} + \psi^p_{it}, \\
q_{it} &= \frac{\phi^s \lambda(T)\varepsilon^d}{\phi^s - \varepsilon^d} T_{it} - \frac{\phi^s \rho^s \varepsilon^d}{\phi^s - \varepsilon^d} Z_{it} + \psi^y_{it},
\end{aligned}$$

and this produces the following partial effects from the reduced form equations:

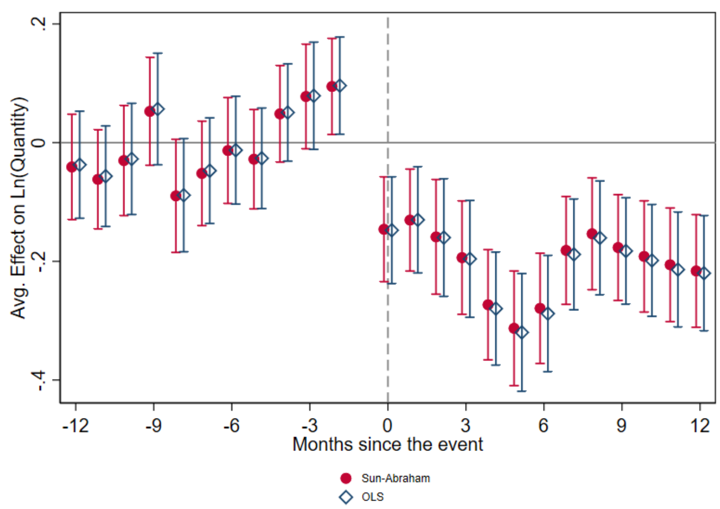
$$\begin{aligned}
\pi_{Tp}(T) &= \frac{\varepsilon^d}{\phi^s - \varepsilon^d} \cdot \left(\frac{d\lambda}{dT} T + \lambda(T) \right), \\
\pi_{Zp} &= -\frac{\phi^s \rho^s}{\phi^s - \varepsilon^d}, \\
\pi_{Tq}(T) &= \frac{\phi^s \varepsilon^d}{\phi^s - \varepsilon^d} \cdot \left(\frac{d\lambda}{dT} T + \lambda(T) \right), \\
\pi_{Zq} &= -\frac{\phi^s \rho^s \varepsilon^d}{\phi^s - \varepsilon^d}.
\end{aligned}$$

Appendix B: Supplemental Tables and Figures

Figure B.1: Event Studies by Treatment Intensity, Quantity



(a) Below Median Treatment Intensity



(b) Above median Treatment Intensity

Notes: Following Callaway et al. (2024), we estimate event studies for groups with different treatment intensities. The figure reports binary event study estimates of the effect of retaliatory tariffs on log quantity for products with below (Panel a) and above (Panel b) median treatment intensity, measured as the added rate as a proportion of the total tariff for each treated product. Each event study includes all untreated varieties as controls. Each panel reports OLS two-way fixed effects and the heterogeneity-robust estimator of Sun and Abraham (2021). Both include variety (country-HS-10) and month-year fixed effects. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Confidence intervals are based on standard errors clustered at the country-HS-6 level.

Table B.1: Out-of-Sample Model Selection

	$K=0$ (Linear)	$K=1$ (Quadratic)	$K=2$ (Cubic)	$K=3$ (Quartic)
Mean OOS MSE	220.214	219.601	219.226	219.421
SD OOS MSE	23.495	23.410	23.320	23.280

Notes: Out-of-sample MSE is computed from 100 repeated random 50/50 splits of the estimation sample. In each split, the IV model is estimated on the training half and evaluated on the held-out test half. Mean and standard deviation (SD) are computed across the 100 splits. K denotes the polynomial order of the compliance function $\lambda(T) = \sum_{k=0}^{K} \theta_k T^k$, which enters the estimating equation as a polynomial of order $K+1$ in T. The sample follows the baseline specification of Table 2, columns (1) and (2).

Table B.2: Instrument Validity: Pre-Period Balance and Sector Decomposition

Panel A: Pre-period Balance Test

	Coefficient	Std. Error	t-stat	p-value
Fossil input share	0.029	0.081	0.35	0.725
Observations		1,957,670		

Panel B: Top 10 sectors by decomposition weight

Sector	Fossil share	w_k	$\hat{\beta}_k$	$w_k \times \hat{\beta}_k$	N
Industrial chemicals	0.052	0.152	0.390	0.059	124,551
Optical & precision instruments	0.004	0.093	−0.318	−0.030	72,070
Electrical machinery	0.005	0.077	0.727	0.056	63,270
Crops & agricultural products	0.053	0.061	−0.481	−0.029	46,393
Stone, cement & minerals	0.075	0.060	0.043	0.003	12,515
Industrial machinery	0.004	0.058	−0.235	−0.014	45,251
Apparel, leather & footwear	0.013	0.056	0.612	0.034	95,989
Iron & steel	0.054	0.055	−0.570	−0.031	38,932
Plastics & synthetic fibers	0.046	0.040	0.341	0.014	55,213
Precision instruments & machinery	0.054	0.032	−0.463	−0.015	21,171
Top 10 total		0.684		0.047	
All 81 sectors		1.000		0.029	

Notes: Panel A reports the coefficient from regressing first-differenced log bilateral export quantities on fossil fuel gross input shares using 2014–2015 data, with country fixed effects and robust standard errors. Panel B decomposes the overall coefficient as $\hat{\beta} = \sum_k w_k \hat{\beta}_k$, where w_k is sector k's share of total within-country instrument variance and $\hat{\beta}_k$ is the sector-specific OLS slope. Weights w_k are non-negative by construction and sum to one.

Table B.3: Tax-Dependent Compliance Rate Estimates (K=1)

	(1)	(2)	(3)	(4)
	Retaliating Countries Only		*All Sample Countries*	
	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)
T	0.320**	-0.941***	0.323**	-0.947***
	(0.119)	(0.218)	(0.119)	(0.214)
T^2	-0.721**	0.963*	-0.726**	0.975*
	(0.244)	(0.548)	(0.244)	(0.540)
Share × ln(Price^f)	4.314***		3.481***	
	(0.617)		(0.419)	
ln(Price)		-1.716***		-1.700***
		(0.183)		(0.151)
Mean	4.145	7.172	4.109	6.762
F-Stat	48.823		69.013	
Observations	1,488,987	1,488,987	3,399,661	3,399,661
Clusters	54,241	54,241	158,672	158,672
$\hat{\theta}_0$		0.548		0.557
$SE(\hat{\theta}_0)$		0.154		0.146
$\hat{\theta}_1$		-0.561		-0.573
$SE(\hat{\theta}_1)$		0.339		0.332
$\hat{\lambda}_1$		0.509		0.517
$SE(\hat{\lambda}_1)$		0.133		0.126
$\hat{\lambda}_2$		0.417		0.422
$SE(\hat{\lambda}_2)$		0.09		0.084
$Avg[\hat{\lambda}_{1i}]$		0.509		0.517
$SE(Avg[\hat{\lambda}_{1i}])$		0.133		0.126
$Avg[\hat{\lambda}_{2i}]$		0.417		0.422
$SE(Avg[\hat{\lambda}_{2i}])$		0.09		0.084
$\tilde{Avg}[\hat{\lambda}_{1i}]$		0.513		0.521
$SE(\tilde{Avg}[\hat{\lambda}_{1i}])$		0.135		0.128
$\tilde{Avg}[\hat{\lambda}_{2i}]$		0.432		0.438
$SE(\tilde{Avg}[\hat{\lambda}_{2i}])$		0.096		0.09
Country-HS-10 FE	x	x	x	x
Country-Month-Year FE	x	x	x	x

Notes: The table reports first-stage, reduced-form, and structural estimates of the tax-dependent compliance function. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped to address potential endogenous tariff targeting. Post-war periods are pooled. Columns (1) and (2) include varieties exported to retaliating countries only; Columns (3) and (4) include all U.S. trading partners. The structural parameters are recovered as described in Section 5. $\hat{\lambda}$ refers to compliance at the mean tax rate among treated observations. $Avg[\hat{\lambda}_i]$ refers to simple average of product-specific compliance rate estimates. $\tilde{Avg}[\hat{\lambda}_i]$ refers to average of product-specific compliance rate estimates weighted by product value. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.4: Tax-Dependent Compliance Rate Estimates (K=3)

	(1)	(2)	(3)	(4)
	Retaliating Countries Only		*All Sample Countries*	
	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)
T	0.424	-2.891***	0.431	-2.897***
	(0.336)	(0.560)	(0.336)	(0.556)
T^2	-1.365	15.322***	-1.394	15.342***
	(2.316)	(4.050)	(2.317)	(4.038)
T^3	1.179	-33.638**	1.229	-33.657**
	(5.536)	(10.314)	(5.537)	(10.291)
T^4	-0.553	22.320**	-0.580	22.329**
	(3.924)	(7.738)	(3.925)	(7.721)
Share × ln(Price^f)	4.311***		3.479***	
	(0.617)		(0.419)	
ln(Price)		-1.708***		-1.695***
		(0.182)		(0.151)
Mean	4.145	7.172	4.109	6.762
F-Stat	48.744		68.947	
Observations	1,488,987	1,488,987	3,399,661	3,399,661
Clusters	54,241	54,241	158,672	158,672
$\hat{\theta}_0$		1.692		1.709
$SE(\hat{\theta}_0)$		0.398		0.379
$\hat{\theta}_1$		-8.971		-9.052
$SE(\hat{\theta}_1)$		2.627		2.564
$\hat{\theta}_2$		19.696		19.859
$SE(\hat{\theta}_2)$		6.461		6.371
$\hat{\theta}_3$		-13.069		-13.175
$SE(\hat{\theta}_3)$		4.770		4.725
$\hat{\lambda}_1$		1.164		1.176
$SE(\hat{\lambda}_1)$		0.257		0.242
$\hat{\lambda}_2$		0.504		0.509
$SE(\hat{\lambda}_2)$		0.100		0.093
$Avg[\hat{\lambda}_{1i}]$		1.219		1.231
$SE(Avg[\hat{\lambda}_{1i}])$		0.271		0.255
$Avg[\hat{\lambda}_{2i}]$		0.579		0.585
$SE(Avg[\hat{\lambda}_{2i}])$		0.112		0.103
$\tilde{Avg}[\hat{\lambda}_{1i}]$		1.302		1.315
$SE(\tilde{Avg}[\hat{\lambda}_{1i}])$		0.293		0.277
$\tilde{Avg}[\hat{\lambda}_{2i}]$		0.679		0.686
$SE(\tilde{Avg}[\hat{\lambda}_{2i}])$		0.134		0.124
Country-HS-10 FE	x	x	x	x
Country-Month-Year FE	x	x	x	x

Notes: The table reports first-stage, reduced-form, and structural estimates of the tax-dependent compliance function. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped to address potential endogenous tariff targeting. Post-war periods are pooled. Columns (1) and (2) include varieties exported to retaliating countries only; Columns (3) and (4) include all U.S. trading partners. The structural parameters are recovered as described in Section 5. $\hat{\lambda}$ refers to compliance at the mean tax rate among treated observations. $Avg[\hat{\lambda}_i]$ refers to simple average of product-specific compliance rate estimates. $\tilde{Avg}[\hat{\lambda}_i]$ refers to average of product-specific compliance rate estimates weighted by product value. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.5: Tax-Dependent Compliance Rate Estimates — Constrained Intercept ($\lambda(0) = 1$)

	(1)	(2)	(3)	(4)
	Retaliating Countries Only		*All Sample Countries*	
	ln(Price) + T	ln(Quantity)	ln(Price) + T	ln(Quantity)
T^2	4.583***	4.787***	4.588***	4.726***
	(0.375)	(0.970)	(0.376)	(0.860)
T^3	-5.092***	-4.460***	-5.099***	-4.392***
	(0.642)	(1.293)	(0.643)	(1.190)
Share × ln(Price^f)	4.390***		3.517***	
	(0.617)		(0.419)	
$\ln(Price) + T$		-1.709***		-1.696***
		(0.179)		(0.150)
Mean	4.179	7.172	4.149	6.762
F-Stat	50.57		70.467	
Observations	1,488,987	1,488,987	3,399,661	3,399,661
Clusters	54,241	54,241	158,672	158,672
$\hat{\theta}_0$ (constrained)		1		1
$\hat{\theta}_1$		-2.801		-2.787
SE($\hat{\theta}_1$)		0.357		0.345
$\hat{\theta}_2$		2.609		2.59
SE($\hat{\theta}_2$)		0.596		0.583
$\hat{\lambda}_1$		0.819		0.82
SE($\hat{\lambda}_1$)		0.022		0.021
$\hat{\lambda}_2$		0.487		0.489
SE($\hat{\lambda}_2$)		0.055		0.053
Avg[$\hat{\lambda}_{1i}$]		0.828		0.829
SE(Avg[$\hat{\lambda}_{1i}$])		0.02		0.019
Avg[$\hat{\lambda}_{2i}$]		0.508		0.51
SE(Avg[$\hat{\lambda}_{2i}$])		0.051		0.05
$\tilde{\text{Avg}}$[$\hat{\lambda}_{1i}$]		0.851		0.852
SE($\tilde{\text{Avg}}$[$\hat{\lambda}_{1i}$])		0.017		0.016
$\tilde{\text{Avg}}$[$\hat{\lambda}_{2i}$]		0.559		0.561
SE($\tilde{\text{Avg}}$[$\hat{\lambda}_{2i}$])		0.046		0.045
Country-HS-10 FE	x	x	x	x
Country-Month-Year FE	x	x	x	x

Notes: The table re-estimates the baseline quadratic compliance function imposing the boundary condition $\lambda(0) = 1$ (full compliance at zero tariff). This constraint is imposed by replacing the endogenous price variable with $\ln(\text{Price}) + T$, where $T = \ln(1 + \tau)$, which absorbs the linear tariff term and fixes $\theta_0 = 1$. The remaining polynomial coefficients θ_1 and θ_2 are estimated freely. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Post-war periods are pooled. Columns (1) and (2) include varieties exported to retaliating countries only; Columns (3) and (4) include all U.S. trading partners. Structural parameters are recovered as described in Section 5. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.6: Tax-Dependent Compliance Rate Estimates - Alternate Fossil Shifters

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Contemporaneous, Share $\times$ $ln(Price_t^f)$				*Lag, Share* $\times$ $(Price_{t-1}^f/Price_{2016}^f)$			
	Retaliating Countries Only		*All Sample Countries*		*Retaliating Countries Only*		*All Sample Countries*	
	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)
T	0.391**	-1.644***	0.396**	-1.643***	0.391**	-1.557***	0.396**	-1.568***
	(0.197)	(0.340)	(0.197)	(0.338)	(0.197)	(0.354)	(0.197)	(0.349)
T^2	-1.089	4.420**	-1.104	4.419**	-1.090	4.177**	-1.105	4.209**
	(0.831)	(1.452)	(0.831)	(1.448)	(0.831)	(1.511)	(0.831)	(1.496)
T^3	0.443	-3.981**	0.455	-3.981**	0.443	-3.876**	0.456	-3.890**
	(0.856)	(1.512)	(0.856)	(1.512)	(0.856)	(1.586)	(0.857)	(1.574)
Share $\times$ ln$(Price_t^f)$	4.773***		3.862***					
	(0.627)		(0.427)					
Share$\times$ $(Price_{t-1}^f/Price_{2016}^f)$					3.108***		2.516***	
					(0.453)		(0.308)	
ln(Price)		-1.519***		-1.520***		-1.726***		-1.699***
		(0.155)		(0.129)		(0.187)		(0.154)
Mean	4.145	7.172	4.109	6.762	4.145	7.172	4.109	6.762
F-Stat	57.918		81.669		47.063		66.857	
Observations	1488987	1488987	3399661	3399661	1488987	1488987	3399661	3399661
Clusters	54241	54241	158672	158672	54241	54241	158672	158672
$\hat{\theta}_0$		1.082		1.08		.902		.923
$SE(\hat{\theta}_0)$		.269		.254		.247		.236
$\hat{\theta}_1$		-2.91		-2.907		-2.42		-2.478
$SE(\hat{\theta}_1)$		1.042		1.012		.955		.938
$\hat{\theta}_2$		2.621		2.618		2.245		2.289
$SE(\hat{\theta}_2)$		1.049		1.032		.969		.962
$\tilde{Avg}[\hat{\lambda}_{1i}]$		.927		.926		.773		.791
$SE(\tilde{Avg}[\hat{\lambda}_{1i}])$		.218		.205		.2		.19
$\tilde{Avg}[\hat{\lambda}_{2i}]$		.62		.619		.521		.533
$SE(\tilde{Avg}[\hat{\lambda}_{2i}])$		.123		.114		.113		.106
Country-HS-10 FE	x	x	x	x	x	x	x	x
Country-Month-Year FE	x	x	x	x	x	x	x	x

Notes: The table reports first-stage, reduced-form, and structural estimates using alternative fossil fuel cost shifters. Columns (1)–(4) use the 2016 industry-level fossil fuel cost share interacted with the contemporaneous log fossil fuel price; Columns (5)–(8) use the 2016 industry-level fossil fuel cost share interacted with the lagged fossil fuel price at time $t-1$ relative to its 2016 level. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Columns (1)–(2) and (5)–(6) include varieties exported to retaliating countries only; Columns (3)–(4) and (7)–(8) include all U.S. trading partners. Structural parameters are recovered as described in Section 5. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.7: Tax-Dependent Compliance Rate Estimates - Robustness to FEs

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Country-HS10 and Month FEs				*Country-HS10, Country-Month, and Industry-Month FEs*			
	Retaliating Countries Only		*All Sample Countries*		*Retaliating Countries Only*		*All Sample Countries*	
	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)
T	0.509**	-1.510***	0.515**	-1.501***	0.294	-1.590***	0.250	-1.667***
	(0.191)	(0.338)	(0.190)	(0.327)	(0.197)	(0.357)	(0.195)	(0.353)
T^2	-1.333*	4.111**	-1.292*	3.916**	-0.782	4.251**	-0.558	4.541**
	(0.768)	(1.396)	(0.765)	(1.359)	(0.829)	(1.518)	(0.826)	(1.515)
T^3	0.665	-3.891**	0.615	-3.650**	0.217	-3.910**	0.006	-4.191**
	(0.790)	(1.475)	(0.787)	(1.443)	(0.851)	(1.600)	(0.847)	(1.607)
Share × $ln(Price^f)$	4.134***		3.316***		3.836***		2.996***	
	(0.609)		(0.415)		(0.862)		(0.578)	
ln(Price)		-1.699***		-1.632***		-1.715***		-1.764***
		(0.188)		(0.154)		(0.293)		(0.249)
Mean	4.145	7.172	4.109	6.762	4.145	7.172	4.109	6.762
F-Stat	46.064		63.985		19.794		26.834	
Observations	1488987	1488987	3399781	3399781	1488987	1488987	3399661	3399661
Clusters	54241	54241	158691	158691	54241	54241	158672	158672
$\hat{\theta}_0$		.888		.919		.927		.944
$SE(\hat{\theta}_0)$		.247		.238		.291		.26
$\hat{\theta}_1$		-2.419		-2.399		-2.478		-2.574
$SE(\hat{\theta}_1)$		.915		.898		1.04		.965
$\hat{\theta}_2$		2.29		2.236		2.28		2.376
$SE(\hat{\theta}_2)$		.93		.925		1.028		.972
$\tilde{Avg}[\hat{\lambda}_{1i}]$		.76		.792		.795		.808
$SE(\tilde{Avg}[\hat{\lambda}_{1i}])$		.202		.193		.24		.212
$\tilde{Avg}[\hat{\lambda}_{2i}]$		.51		.543		.536		.539
$SE(\tilde{Avg}[\hat{\lambda}_{2i}])$		.116		.11		.143		.125
Country-HS-10 FE	x	x	x	x	x	x	x	x
Country-Month-Year FE					x	x	x	x
Month-Year FE	x	x	x	x				
Industry-Month-Year FE					x	x	x	x

Notes: The table reports first-stage, reduced-form, and structural estimates under alternative fixed effects specifications. Columns (1)–(4) replace the baseline country-month-year fixed effects with month-year fixed effects, providing a less demanding specification. Columns (5)–(8) augment the baseline with industry-month-year fixed effects, which absorb demand shocks common to products within 15 broad industry categories and directly address the concern that energy price variation may proxy for sector-specific demand conditions. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Post-war periods are pooled. Columns (1)–(2) and (5)–(6) include varieties exported to retaliating countries only; Columns (3)–(4) and (7)–(8) include all U.S. trading partners. Structural parameters are recovered as described in Section 5. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.8: Tax-Dependent Compliance Rate Estimates
(Excluding High Value Products)

	(1)	(2)	(3)	(4)
	Retaliating Countries Only		*All Sample Countries*	
	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)
T	0.365*	-1.685***	0.369*	-1.686***
	(0.203)	(0.365)	(0.203)	(0.363)
T^2	-1.013	4.793**	-1.029	4.796**
	(0.863)	(1.556)	(0.863)	(1.551)
T^3	0.368	-4.354**	0.381	-4.355**
	(0.878)	(1.600)	(0.879)	(1.599)
Share × ln($Price^f$)	4.351***		3.496***	
	(0.619)		(0.420)	
ln(Price)		-1.682***		-1.679***
		(0.179)		(0.150)
Mean	4.149	7.161	4.11	6.757
F-Stat	49.345		69.398	
Observations	1482553	1482553	3393227	3393227
Clusters	54211	54211	158642	158642
$\hat{\theta}_0$		1.001		1.003
$SE(\hat{\theta}_0)$		.26		.247
$\hat{\theta}_1$		-2.849		-2.855
$SE(\hat{\theta}_1)$		1.015		.986
$\hat{\theta}_2$		2.588		2.593
$SE(\hat{\theta}_2)$		1.008		.991
$\tilde{Avg}[\hat{\lambda}_{1i}]$		.835		.837
$SE(\tilde{Avg}[\hat{\lambda}_{1i}])$		.205		.194
$\tilde{Avg}[\hat{\lambda}_{2i}]$		.566		.567
$SE(\tilde{Avg}[\hat{\lambda}_{2i}])$		.121		.113
Country-HS-10 FE	x	x	x	x
Country-Month-Year FE	x	x	x	x

Notes: The table reports first-stage, reduced-form, and structural estimates applying a broader exclusion restriction than the baseline. While the baseline drops only treated HS-10 products that accounted for all HS-6 category value in 2017, this specification drops all treated HS-10 products whose 2017 export share within their destination–HS-6 group exceeds the 25th percentile of the share distribution, applied to treated products. This provides a robustness check on whether the baseline results are sensitive to the choice of exclusion threshold for addressing potential endogenous tariff targeting. Post-war periods are pooled. Columns (1) and (2) include varieties exported to retaliating countries only; Columns (3) and (4) include all U.S. trading partners. Structural parameters are recovered as described in Section 5. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.9: Country Heterogeneity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	EU + Canada				*CH, MX, RU, TR*			
	ln(price)	ln(Quantity)	ln(price)	ln(Quantity)	ln(price)	ln(Quantity)	ln(price)	ln(Quantity)
T	0.176	-1.860***	0.190	-1.862***	0.149	-1.280**	0.151	-1.284**
	(0.295)	(0.430)	(0.295)	(0.428)	(0.221)	(0.416)	(0.222)	(0.415)
T^2	0.543	5.406***	0.495	5.403***	-0.805	2.106*	-0.803	2.122*
	(1.034)	(1.544)	(1.035)	(1.541)	(0.588)	(1.215)	(0.588)	(1.209)
T^3	-5.674***	-15.388***	-5.625***	-15.346***	0.553	-1.841	0.547	-1.851
	(1.324)	(3.077)	(1.326)	(3.020)	(0.673)	(1.375)	(0.673)	(1.368)
Share × ($Price^f$)	4.213***		3.412***		4.395***		3.512***	
	(0.637)		(0.425)		(0.624)		(0.421)	
ln(Price)		-1.694***		-1.686***		-1.725***		-1.705***
		(0.189)		(0.155)		(0.181)		(0.151)
Mean	4.153	7.131	4.111	6.736	4.157	7.159	4.114	6.754
F-Stat	43.794		64.481		49.547		69.538	
Observations	1412253	1412253	3322927	3322927	1462512	1462512	3373186	3373186
Clusters	52927	52927	157358	157358	53642	53642	158073	158073
$\hat{\theta}_0$		1.097		1.103		.742		.752
$SE(\hat{\theta}_0)$		.294		.282		.258		.255
$\hat{\theta}_1$		-3.19		-3.204		-1.22		-1.244
$SE(\hat{\theta}_1)$		.968		.952		.731		.728
$\hat{\theta}_2$		9.083		9.1		1.067		1.085
$SE(\hat{\theta}_2)$		1.766		1.757		.813		.814
$\tilde{Avg}[\hat{\lambda}_{1i}]$		1.088		1.094		.663		.672
$SE(\tilde{Avg}[\hat{\lambda}_{1i}])$		.291		.279		.215		.213
$\tilde{Avg}[\hat{\lambda}_{2i}]$		.869		.874		.539		.545
$SE(\tilde{Avg}[\hat{\lambda}_{2i}])$		.191		.18		.152		.151
Country-HS-10 FE	x	x	x	x	x	x	x	x
Country-Month-Year FE	x	x	x	x	x	x	x	x

Notes: The table reports first-stage, reduced-form, and structural estimates separately for subgroups of retaliating countries. Columns (1)–(4) includes estimates for the EU and Canada, which represent economies with strong customs enforcement institutions; China, Mexico, Russia, and Turkey (Columns (5)–(8)) represent economies with weaker enforcement. The comparison across groups tests whether compliance patterns differ systematically with institutional quality, with similar patterns suggesting that legal avoidance rather than illegal evasion drives the results. Columns (1)–(2) and (5)–(6) include only the treated countries in the column heading plus untreated products from retaliating countries; Columns (3)–(4) and (7)–(8) additionally include untreated products from all non-retaliating control countries. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Post-war periods are pooled. Structural parameters are recovered as described in Section 5. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.10: Effects of Added Rates on Untreated Units

	(1)	(2)	(3)	(4)	(5)	(6)
	Retaliating Countries Only			*All Sample Countries*		
	ln(Quantity)	ln(Quantity)	ln(Quantity)	ln(Quantity)	ln(Quantity)	ln(Quantity)
1[Any Added Rate in HS4]	0.018			0.017		
	(0.019)			(0.019)		
ln(1+ Max Added Rate in HS4)		0.116			0.112	
		(0.098)			(0.097)	
ln(1 + Mean Added Rate in HS4)			0.300			0.313
			(0.252)			(0.250)
Observations	1385778	1385778	1385778	3296452	3296452	3296452
Clusters	52328	52328	52328	156759	156759	156759
Country-HS10 FE	x	x	x	x	x	x
Country-Month-Year FE	x	x	x	x	x	x

Notes: The table reports estimates of spillover effects of retaliatory tariffs on untreated HS-10 products within the same HS-4 category and country. Positive coefficients are consistent with either substitution toward untreated varieties or misclassification of treated products into untreated HS codes to avoid tariffs. The sample includes only untreated HS-10 products. Tariff exposure measures vary at the HS-4-country-month level; Columns (1) and (4) use an indicator for any positive tariff within the HS-4 category, Columns (2) and (5) use the maximum added rate, and Columns (3) and (6) use the mean added rate. Columns (1)–(3) include varieties exported to retaliating countries only; Columns (4)–(6) include all U.S. trading partners. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level are in parentheses. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table B.11: Estimating λ with Misclassification

	(1)	(2)	(3)	(4)
	Retaliating Countries Only		*All Sample Countries*	
	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)
T	0.350**	-1.305***	0.355**	-1.312***
	(0.167)	(0.298)	(0.167)	(0.295)
T^2	-0.942*	2.349**	-0.952*	2.362**
	(0.513)	(0.983)	(0.513)	(0.974)
T^3	0.287	-2.312**	0.296	-2.318**
	(0.606)	(1.168)	(0.606)	(1.163)
ln(1 + Mean Added Rate in HS4)	0.075	0.043	0.074	0.046
	(0.088)	(0.135)	(0.088)	(0.134)
Share × $(Price^f)$	4.308***		3.478***	
	(0.617)		(0.419)	
$ln(Price)$		-1.713***		-1.698***
		(0.183)		(0.151)
Mean	4.145	7.172	4.109	6.762
F-Stat	48.68		68.896	
Observations	1488987	1488987	3399661	3399661
Clusters	54241	54241	158672	158672
$\hat{\theta}_0$		.762		.772
$SE(\hat{\theta}_0)$		.209		.199
$\hat{\theta}_1$		-1.371		-1.391
$SE(\hat{\theta}_1)$		.621		.607
$\hat{\theta}_2$		1.349		1.365
$SE(\hat{\theta}_2)$		.706		.702
$\tilde{Avg}[\hat{\lambda}_{1i}]$		.69		.699
$SE(\tilde{Avg}[\hat{\lambda}_{1i}])$		.179		.169
$\tilde{Avg}[\hat{\lambda}_{2i}]$		.55		.558
$SE(\tilde{Avg}[\hat{\lambda}_{2i}])$		.123		.116
Country-HS10 FE	x	x	x	x
Country-Month-Year FE	x	x	x	x

Notes: The table reports first-stage, reduced-form, and structural estimates controlling for the mean tariff rate within the surrounding HS-4 category (excluding the product's own HS-6) to assess whether misclassification of treated products into nearby untreated codes biases the baseline compliance estimates. If misclassification were driving the results, this control should partially absorb the bias and alter the compliance estimates. The sample is fixed to include all country-HS-10 products observed throughout the entire sample period. Treated HS-10 products that accounted for all HS-6 category value in 2017 are dropped. Post-war periods are pooled. Columns (1) and (2) include varieties exported to retaliating countries only; Columns (3) and (4) include all U.S. trading partners. Structural parameters are recovered as described in Section 5. All specifications include country-HS-10 and country-month-year fixed effects. Standard errors clustered at the country-HS-6 level. *** p<0.01, ** p<0.05, * p<0.1.