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 nearly 30 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 accounts for approximately $6 billion in foregone tariff revenue.

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

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

1 Introduction

Import tariffs are a pillar-stone economic policy in the strategic interaction among countries. The recent trade war reminds economists and policymakers of the influence of tariffs, which seemed to lost salience since the establishment of the WTO. To understand the economic effects of tariffs, we need to know the effectively implemented tariff rates. This work develops an innovative method for identifying and estimating tariff compliance rates.

Just like any form of taxes, buyers and sellers in international trade have the incentive to evade tariffs in order to reduce costs. Recent evidence from the U.S. Customs and Border Protection shows that during the 2018 trade war, enforcement actions under the Enforce and Protect Act (EAPA) rose sharply, with the value of investigated cases increasing from $15 million in 2018 to $215 million in 2020 (Roll and Akers, 2023).

Noncompliance with tariffs can take many forms. At the border customs, to evade tax payments, importers can underreport quantities, underreport unit prices, or misclassify a product from a high-tariff code to a low-tariff code. In the long run, firms may also re-route their shipments through a third country facing lower tariffs. This makes tariff evasion hard to observe and measure. In addition, the evasion of tariffs increases with tariff rates, as found by Fisman and Wei (2004). Last, governments can exempt specific firms from import duties and strategically adjust free-trade regulations when they must increase tariffs. All of these add to the difficulties of knowing the effectively implemented tariff rates.

In this work, we propose an innovative identification method for short-run tariff compliance rates by comparing the effects of exogenous tax variation with the effects of exogenous variation in tax-exclusive prices. Specifically, in a two-period model with perfect competition, using a supply shifter as an instrument for tax-exclusive prices enables us to identify and estimate the demand elasticity. Meanwhile, compliance rates are identified by how quantities respond to changes in supply shifters and tariff rates. Our method allows the compliance rates to be either constant or tax-dependent.[1]

[1]We show in Appendix A.2 that this identification method also holds under imperfect competition.

It is worth noting that the short-run noncompliance of tariffs identified by our method includes both the evaded and avoided shares in the statutory tariff rates, achieved through all possible means. In other words, we can accurately derive the effectively implemented tariff rates.[2]

We apply this method to six countries that retaliated against US tariffs during the 2018 trade war.[3] We build a panel dataset with monthly U.S. export values by products and importing countries during 2017-2019. Event studies show an immediate and persistent decline in U.S. export quantities following increases in the tariffs they face. A 10% increase in the tariff rate by a partner country decreases U.S. export quantities by 8%, while U.S. export prices do not change significantly.

Using fossil fuel costs in production as the supply shifter, we estimate the average demand elasticity for U.S. exports as -1.7. Compliance rates among the six U.S. trade partners decreased from approximately 80% before the trade war in 2017 to approximately 50% after the tariff increases in 2018 and 2019, with average product-level statutory tariffs rising from 7% to 26% over the period. We also find that compliance rates decrease with tariff rates and that a 1 percentage point increase in the tariff rate reduces compliance by more than 1 percentage point. Our results are robust under various specifications and sample selections.

Using the parameter estimates, we calculate that, after the 2018 trade war, the overall average effectively implemented tariffs faced by U.S. exports to the six retaliatory trade partner countries increased by only 0.1% for intermediate goods and 0.25% for final goods. Meanwhile, the decrease in compliance rates during the trade war generated $6B tariff revenue loss for the six U.S. trade partners in 2019. These results highlight the value of estimating tariff compliance during the trade war.

Related Literature. Our work contributes to three strands of literature: estimating

[2]Our identified noncompliance of tariffs is short-rum because re-routing shipments are not included.

[3]These include China, Mexico, Turkey, the European Union, Canada, and Russia. While India imposed retaliatory tariffs on U.S. products in 2019, we focus on the additional tariffs imposed in 2018, which provide sufficient data for the period after the tariff changes.

the level of tariff evasion, analyzing the effects of trade wars, and estimating tax compliance more broadly. The level of tariff compliance is an open empirical question. The literature on imperfect tariff compliance begins with Fisman and Wei (2004), who estimate changes in tariff evasion using gaps between reported trade values reported by importers and exporters. They find that a 1 percentage point increase in the tax rate increases trade value gaps by 3%. The works of Mishra et al. (2008), Javorcik and Narciso (2017), Beverelli and Ticku (2022), and Kitsios et al. (2023) apply this method to study various factors affecting tariff evasion, including WTO regulations, trade facilitation, and customs enforcement. While this approach captures changes in evasion relative to tariff adjustments induced by factor shocks, it does not reveal the overall level of compliance. Our work contributes to this literature by being the first study to provide a direct estimate of the level of tariff compliance. Estimating the compliance rate function also allows us to calculate the marginal effects of tax increases. Our results imply that a one percentage point increase in the tariff rate increased evasion by 2 to 10% in our sample, which is in line with the 3% marginal effects found by Fisman and Wei (2004).

Our empirical application also relates to the nascent literature on the recent U.S. trade wars. Fajgelbaum et al. (2020), Flaaen et al. (2020b), and Cavallo et al. (2021) find that U.S. import tariffs have complete pass-through to importer prices, while evidence on the pass-through to exporter prices is mixed. We bring new insights to the estimation of trade elasticities when countries deviate from free trade. We estimate a non-trivial level of evasion by foreign importers of U.S. goods; this implies that ignoring non-compliance and inferring importer prices from exporter prices and the statutory tariff rates, as considered in previous literature, leads to an underestimation of demand elasticities and an overestimation of pass-through to importer prices for U.S. exports.

Finally, our approach to estimating tariff-dependent compliance rates from variation in tariffs extends the literature on using variation in tax rates to identify market parameters and on estimating tax compliance more broadly. The approach can be applied in a variety

of settings in which compliance with sales and excise taxes is unknown. Romalis (2007) and Zoutman et al. (2018) reveal how variation in tax rates identifies both demand and supply elasticities in a perfectly competitive market. Dearing (2022) extends this to imperfect competition. Bibler et al. (2024) extend this technique to estimate tax compliance using changes in tax enforcement, showing that an additional shifter allows for the estimation of tax compliance along with both elasticities. Our work differs from these in two important ways. First, our approach relies on variation in tax rates, which is a relatively common occurrence, rather than changes in tax enforcement. Second, we allow compliance to vary with the tax rate. Thus, our approach extends the number of cases for which tax compliance estimation is feasible.

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, Russia, and India.

We analyze the effects of these retaliatory tariffs imposed on U.S. exports. To do so, we 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.[4]

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

[4]Compared to 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.

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 $122 billion in U.S. annual exports, representing 11.2% of total U.S. exports in 2017 and covering 7,279 product varieties. Columns 6–8 report the average tariffs on U.S. exports, which increased significantly during this period. The trade-weighted average tariff increased from 5.7% to 10.5% from 2017 to 2018, and further increased to 19.7% by 2019. The table highlights significant cross-sectional and temporal variation that we exploit in our empirical application.

Figure 1 illustrates the evolution of retaliatory tariffs from January 2018 to March 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-country settings (Ferrantino et al., 2012; Stoyanov, 2012), developing and transition 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, they can be sizable in the aggregate: Kitsios et al. (2023) reports 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, including exemptions, duty-drawback and processing-trade regimes, bonded warehouses, and Free Trade Zones (FTZs), all of which can reduce or eliminate tariff liability for eligible transactions. These regimes are often designed to facilitate the import of inputs, making intermediate goods more likely to qualify for duty-free treatment (Siroën et al., 2014; Grant, 2020).

The 2018–2019 trade war provides an environment in which both incentives for illegal evasion and opportunities for legal avoidance likely changed sharply. On the U.S. import side, Che et al. (2025) document increased tariff evasion during the trade war, with stronger responses among intermediate goods and among importers facing larger tariff liabilities. A related adjustment margin is re-routing through third countries: using firm-level evidence, Iyoha et al. (2024) estimate that re-routing of Chinese exports to the United States via Vietnam accounted for 1.8% of Vietnamese exports to the United States in 2021, suggesting that re-routing is modest relative to broader trade diversion in that setting.

Much less is known about compliance and avoidance in countries imposing retaliatory tariffs, largely due to limited access to administrative customs data. Nevertheless, evidence suggests that retaliating countries shaped the effective incidence of tariff policy through both product targeting and administrative regimes. Using satellite readings on night-time luminosity, Chor and Li (2024) argue that China's import tariffs on inputs did not generate systematic adverse effects on aggregate activity, in part because China targeted products with substantial scope for trade diversion (i.e., readily available alternative source countries) and in part because it reduced administrative barriers for firms to engage in processing trade

beginning on January 1, 2019.

Taken together, this evidence indicates that statutory tariffs may translate imperfectly into effective tariff burdens. Less than full compliance can reflect illegal evasion at the border, through underreporting, misclassification, and related practices, and legal avoidance through institutional arrangements such as exemptions, duty-free regimes, and special customs procedures. Because many of these channels operate through the same observed trade aggregates, and because direct measures of tariff payments are typically unavailable at the relevant level of disaggregation, we treat observed responses to tariff changes as reflecting the net effect of both evasion and avoidance 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. (2024), 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:

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

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:

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

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 $\rho^d = 0$.[5]

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 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:

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

[5] As outlined by Bibler et al. (2024) and shown in Appendix A.1, whether a supply or demand shifter is needed depends on the statutory burden of the tax.

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 $\pi_{Tp} < 0$. Therefore, the price response enters (4) with a minus sign. The compliance rate captures the extent to which the tax-induced demand shift translates into effective price and quantity adjustments.

Considering 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 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:

$$\begin{aligned} \pi_{Tq}(T) &= \frac{\varepsilon^d \varepsilon^s}{\varepsilon^s - \varepsilon^d}\big[\lambda'(T;\Theta)T + \lambda(T;\Theta)\big], \qquad \pi_{Tp}(T) = \frac{\varepsilon^d}{\varepsilon^s - \varepsilon^d}\big[\lambda'(T;\Theta)T + \lambda(T;\Theta)\big], \\ \pi_{Zq} &= \frac{-\rho^s \varepsilon^d}{\varepsilon^s - \varepsilon^d}, \qquad\qquad\qquad\qquad\qquad\qquad \pi_{Zp} = \frac{-\rho^s}{\varepsilon^s - \varepsilon^d}. \end{aligned}$$

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 first-order differential 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.[6]

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 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 declines with the tax level. 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

[6]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 Θ.

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 the 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.

Effective implementation may depart from full compliance through several institutional 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. Moreover, legal avoidance can arise through duty-free regimes, exemptions, and related institutional features that reduce or eliminate tariff liability for specific transactions. For this reason, in our application, we interpret λ as an *effective compliance (implementation) rate*, that is, the fraction of the statutory tariff that is reflected in observed equilibrium responses under our maintained assumptions, irrespective of whether the underlying mechanism is evasion or 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., 2020a). 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 significant 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.

Finally, our empirical analysis uses multiple goods (or product varieties). In Appendix A, Zoutman et al. (2018) shows 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 the additional retaliatory increment that takes effect. Tariffs are measured at the six-digit Harmonized System (HS6) level.[7]

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 enacted tariff rates. For example, Mexico and Canada imposed zero tariffs on U.S. exports in 2017 under the North American Free Trade Agreement. Tariffs in 2018–January 2020 are calculated as the sum of the 2017 implemented tariffs and the additional retaliatory measures. Notably, this approach isolates statutory policy variation and does not capture potential trade flow adjustments in response to changes in tariff rates.

The monthly administrative U.S. export records are from the U.S. Census Bureau, which

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

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 the largest trade flows most likely to shape tariff-setting, that is, the most-traded narrow product line within each destination–HS6 cell.[8]

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. 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, direct fossil fuel input shares are 1.2 percent, while gross shares are 3.2 percent, implying that I–O linkages roughly triple sectors' exposure to fossil fuel price changes.[9] 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.

[8]Specifically, we sum monthly exports over 2017 to compute total pre-war export value for each destination–HS10 pair, rank HS10 products within each destination–HS6 group, and exclude products with rank below two. This procedure drops the unique highest-value HS10 product; in the case of a two-way tie for first, both products are dropped (average rank 1.5), while in the case of a three-way tie no product is dropped (average rank 2). We apply this exclusion only within treated destination–HS6 groups.

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

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 trade value (if $y = v$), price (if $y = p$), or quantity (if $y = q$) of variety i in month t. A variety is defined as an importer-HS10 product pair. 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.

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. Because both variety and time fixed effects are included, the β_j coefficients are identified from differential within-variety changes over time relative to never-treated varieties.

Figures 3, 4, and 5 display the dynamic responses of U.S. export values, quantities, and prices around the introduction of retaliatory tariffs. Similar to Fajgelbaum et al. (2020) (Figure III), there is no clear evidence of anticipatory behavior prior to the imposition of tariffs: the estimated coefficients for the twelve months preceding the event fluctuate around zero.

Following the imposition of retaliatory tariffs ($j = 0$), export values and quantities exhibit large and sustained declines (Figures 3). Within six months of tariff imposition, export volumes and values fall by approximately 20 to 30 percent relative to pre-treatment levels, suggesting that adjustment occurs primarily through the intensive margin rather than com-

plete price adjustment. The similarity between the standard two-way fixed effects (TWFE) and the Sun-Abraham estimators indicates that heterogeneity in treatment timing does not drive the results. The Poisson specification in Figure 4 further confirms these findings in a framework that accommodates zero trade flows. The post-treatment dynamics are nearly identical to those from the linear models, with slightly larger effects in magnitude. The magnitude of these declines is similar to Fajgelbaum et al. (2020), who find that at the month of implementation, export values decline on average by 24% and quantities fall by 25%.

The average post-event effect on export prices is small and statistically insignificant (Figure 5). Retaliatory tariffs do not lead to measurable U.S. exporter price reductions. Our results are consistent with broader evidence. Amiti et al. (2019) (Table 3) find no decline in US 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.

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 first- and second-stage equations:

$$\begin{aligned} p_{it} &= \pi_{Tp} T_{it} + \pi_{Zp} Z_{it} + \delta_i + \delta_{ct} + u^p_{it}, \\ q_{it} &= \varepsilon^d p_{it} + \gamma^d T_{it} + \rho^d Z_{it} + \delta_i + \delta_{ct} + u^d_{it}, \end{aligned}$$

where i indexes varieties and t indexes months. 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 retaliatory 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 3.1, 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 3.1. 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,

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

so that the estimating equation becomes

$$q_{it} = \varepsilon^d p_{it} + \varepsilon^d \sum_{k=0}^{K} \theta_k T_{it}^{k+1} + \delta_i + \delta_{ct} + u^d_{it}. \tag{8}$$

5.2 Model Selection and Baseline Specification

To determine the optimal polynomial order K for the compliance function $\lambda(T)$, we employ out-of-sample (OOS) cross-validation. We estimate the model for $K = 0, \ldots, 3$ on a randomly selected training subset of the data and evaluate its predictive accuracy on the held-out test set. Table A1 reports the mean squared errors (MSE) across specifications and indicates that

a quadratic specification ($K = 2$) provides the best fit, though higher-order polynomials yield similar performance. We thus adopt the quadratic form for $\lambda(T)$:

$$\lambda(T) = \theta_0 + \theta_1 T + \theta_2 T^2. \tag{9}$$

Substituting into the demand equation, and using $\pi_{Tp} = \tilde{\pi}_{Tp}\lambda(T)$, yields the following cubic specification in the tariff variable T for our baseline first-step estimation:

$$p_{it} = \tilde{\pi}_{Tp}\theta_0 T_{it} + \tilde{\pi}_{Tp}\theta_1 T_{it}^2 + \tilde{\pi}_{Tp}\theta_2 T_{it}^3 + \pi_{Zp} Z_{it} + \delta_i + \delta_{ct} + u_{it}^p. \tag{10}$$

The baseline second-step estimation is then:

$$q_{it} = \varepsilon^d p_{it} + \varepsilon^d \theta_0 T_{it} + \varepsilon^d \theta_1 T_{it}^2 + \varepsilon^d \theta_2 T_{it}^3 + \delta_i + \delta_{ct} + u_{it}^d. \tag{11}$$

Our estimation proceeds via two-stage least squares, instrumenting p_{it} with the supply shifter Z_{it}. We assume the higher-order terms in T_{it} are exogenous. 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 9:

$$\frac{d\lambda(T)}{d\tau} = \frac{\theta_1}{1+\tau} + \frac{2\theta_2 \ln(1+\tau)}{1+\tau}. \tag{12}$$

Identification and Interpretation 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–month fixed effects, the

shifter satisfies the exclusion restriction (SER2) in Section 3.1. Its role is analogous to that of the tax instrument in the structural estimation strategy of Zoutman et al. (2018): it generates exogenous variation in equilibrium prices that identifies the demand elasticity and the compliance function.[10]

A potential concern is that energy prices may proxy for broader global business-cycle 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. Remaining threats 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}. We address this concern in two ways. First, we show that the results are robust to alternative constructions of the energy-cost shifter, including using contemporaneous rather than lagged fossil fuel prices and an alternative normalization of the fossil fuel price series (Table B.1). Second, we verify that the estimates are stable to richer fixed-effect structures that absorb broader sectoral demand shocks, including specifications with month-year fixed effects and with industry–month-year fixed effects (Table B.2).

6 Estimation Results

In this section, we discuss estimation results for the first stage and demand equations, as well as tariff compliance rates before and after the trade war. In addition, we estimate differential rates for intermediate goods and final goods, and for differentiated and non-differentiated products.

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

6.1 Baseline Results

Table 2 shows baseline estimates of Equations (10) and (11) in Section 5. Columns (1) and (2) show results using observations of retaliating countries. Columns (3) and (4) show results using observations of all US trading partners. In the baseline estimation, we include country-HS10 product fixed effects and country-month-year fixed effects.

Column (1) in Table 2 shows a statistically significant relation between the supply shifter and the natural log of prices. The F-stat is 52, implying a strong first stage which is necessary for our identification strategy. Meanwhile, none of the three coefficients for the tariff terms are statistically significant, which implies a flat supply function. In Column (2) of Table 2, the coefficient on $ln(price)$ indicates a demand elasticity of -1.7. The ratios between the three coefficients of the tariff terms and the demand elasticity reveal the three components in the compliance rate of tariff, which are θ_0, θ_1, and θ_2 from Equation (11) in Section 5. Specifically, when using observations of retaliating countries only, the estimated compliance function is $\widehat{\lambda(T)} = 0.97 - 2.4T + 2.1T^2$.

We calculate the pre-war and post-war tariff compliance rates according to the estimated relation between compliance and tariff rates, which are displayed near the bottom of Table 2. To obtain the estimated pre-war compliance rate, $\hat{\lambda}_1$, we use the average tariffs for retaliating varieties during pre-war months, $\overline{ln(1+\tau_{ict})} \approx 0.07$. The pre-war compliance rate is approximately 0.81 (81%) with a standard error of 0.2, which was obtained using the delta method. We estimate a pre-war marginal effect of increasing tariffs on compliance of -1.9.[11] That is, a 1 percentage point increase in the tariff rate decreases the rate of compliance by 1.9 percentage points. In other words, we find that, at pre-war tariff rates, a 1 percentage point increase in tariffs leads to a 10% increase in evasion.

To obtain the post-war compliance rate, $\hat{\lambda}_2$, we use the average T for retaliating varieties during months after the tariff increases were initiated, or $\overline{ln(1+\tau_{ict})} \approx 0.235$. The estimated post-war compliance rate is approximately 0.52 (52%) with a standard error of 0.1. The post-

[11] At the mean tariff rate among treated products in the pre-treatment period.

war marginal effect of a tariff rate increase on compliance is -1.1. That is, a 1 percentage point increase in the tariff rate decreases compliance by 1.1 percentage points.[12] In terms of evasion, a 1 percentage point increase in tariff rates leads to a 2.3% increase in evasion at post-treatment tariff levels. Thus, the results reveal that evasion increases by 2.3-10% with every 1 percentage point increase in tariffs, which aligns with the 3% increase found by Fisman and Wei (2004).[13]

There are two important takeaways from the relationship between compliance and tariff rates. First, we find significant decreases in the level of compliance during the trade war, which is apparent from the change in compliance in the period after the higher tariffs were implemented. This is consistent with findings in previous literature following Fisman and Wei (2004). Second, we can derive the marginal effect of tariff changes on the compliance rate following Equation (12) in Section 5, which allows for a more direct comparison with prior literature.

Robustness: To test the robustness of our baseline results, we include estimates based on alternative cost shifter calculations and alternative fixed effects specifications. In Table B.1, we use two alternative cost shifters. In columns (1)-(4) of Table B.1, we replace the first lag of fossil prices by the contemporaneous fossil price in the supply shifter. The results are close to the baseline estimates. The estimated demand elasticity is approximately -1.5. The pre-war compliance is between 79% and 93%. The post-war compliance is between 51% and 59.7%. In columns (5)-(8) of Table B.1, we apply an alternative measure of the fossil fuel prices, using the ratio of the price at time t and the 2016 price. The results are also close to our baseline results. Using this alternative shifter calculation, we find pre-war compliance rates of about 85%, and post-treatment compliance rates around 55%.

[12] At the mean tariff rate among treated products in months with the higher tariffs in place.

[13] In Columns (3) and (4) of Table 2, we show the analogous results when using observations for all U.S. trade partners. The estimates of demand elasticity and tariff terms are similar to the baseline results in Columns (1) and (2). The estimated pre-war compliance rate is 0.71 with a standard error of 0.144, and the estimated post-war compliance rate is 0.46 with a standard error of 0.08.

Table B.2 show results under two alternative estimation specifications. In columns (1)-(4), we include country-HS10 product fixed effects and month-year fixed effects in the estimation. In columns (5)-(8), we include three sets of fixed effects: Country-HS10 product fixed effects, Country-month-year fixed effects, an industry-month-year fixed effects.[14] In all cases, the results are comparable to the main results. Estimated demand elasticities range from -1.69 to -1.87. The pre-war compliance rates are between 68% and 85%, and the post-war compliance rates are between 44% and 53%. Overall, the results are robust to alternative cost shifters and changes to the empirical specification.

6.2 Intermediate and Final Goods

Intermediate goods are eligible for duty-free importing in most countries. Duty-free imports can take several different forms, such as processed trade, goods shipped to bounded warehouse and free trade zones. To separate the effect of duty-free imports from tariff evasion, we categorize products into intermediate goods and final, consumption goods using the Broad Economic Categories (BEC) system. In the estimation, we use the interaction terms between explanatory variables and product type indicators. This allows us to estimate separate demand elasticities and compliance rates for each product type. The results are presented in Table 3.

In column (1) of Table 3, we present a single first stage equation. There is a strong first stage relationship, an F-stat of 77.4, which supports our use of the cost shifter as an instrument for prices. From Column (2), the estimated demand elasticity for final goods is -3.3, and the estimated demand elasticity for intermediate goods is -1.7. As in the baseline results, the three components (θ_0, θ_1, and θ_2) in the compliance rate function are estimated using ratios between the coefficients on the tariff terms and the demand elasticities. We present the estimated parameters of the compliance function and the compliance rates in Table 4. The zero-order component of the compliance rate function, θ_0, is lower for inter-

[14] Using 15 broad industry categories.

mediate goods than for final goods, meaning that we estimate higher levels of compliance among final goods when tariff rates are near zero or relatively low. This result highlights the fact that intermediate goods are more likely to be eligible for duty-free importing than final goods. The signs of the first-order and second-order component of compliance rates, θ_1 and θ_2, indicate that compliance rates are decreasing in tariffs for both product types over the relevant range that we study here.

With the estimated relations between compliance and tariff rates, we calculate the pre-war and post-war compliance rates. In Table 4, the average pre-war compliance rate for final goods is 0.7, with a standard error of 0.33. As for the marginal effect, a one percentage point increase in the tariff rate reduces compliance by 1 percentage point. The post-war compliance rate for final goods is 0.37, with a standard error of 0.16, and a one percentage point increase in the tariff rate reduces compliance rate by 2 percentage points. The results suggest that the trade war tariffs greatly decreased the compliance for final goods. Meanwhile, the tariff evasion activities by final-good importers are marginally more responsive to tariff increases in post-war periods than in the pre-war periods.

The average pre-war compliance rate for intermediate goods is around 0.1 lower than that of final goods, and the post-war compliance rates for the two types of products are similar. The standard errors on the estimated compliance rates for intermediate goods are smaller than the corresponding standard errors for the estimated compliance rates among final goods. Retaliating products include more intermediate goods than final goods, with about 75% of the observations falling in the intermediate goods category. The marginal effects of increasing tariffs on compliance for intermediate goods are also comparable to those of final goods. A one percentage point increase in the tariff rate reduces the compliance rate for intermediate goods by 1.7 and 1 percentage points in the pre- and post-war periods, respectively. These results show that the trade war tariffs also largely decrease the compliance for intermediate goods. However, the tariff evasion activities by intermediate-good importers are marginally less responsive to tariff increases in post-war periods than in the pre-war periods.

The estimated compliance rate generate implications about effective tariff rates and government revenue loss from non-compliance, as shown in Table 7. In 2017, before the trade war, the U.S. exported a total of $76.7 billions worth of final goods to the retaliating countries. The average statutory tariffs faced for U.S. final good exports was 1.36%. Our results suggest the average effective tariff rate was only 0.95%, when taking non-compliance into consideration. The tariff revenue loss for governments of the retaliating countries was around $0.3 billion from final good imports. As for the intermediate goods, the average statutory tariff rate was 1.8% and the compliant effective tariff rate was around 1.16%, on average. With total U.S. exports of $662 billion to the retaliating countries, the tariff revenue loss for the governments in the retaliating countries was around $4.4 billion in intermediate good imports. It is worth noting that this $4.4 billion partially comes from duty-free imports.

In 2019, after the trade war-induced tariff hikes were in place, the average statutory tariff applied to U.S. final good exports increased to 3.27%. With a total U.S. export of $80.3 billion to the retaliating countries, the tariff revenue loss for the governments in the retaliating countries was around $1.66 billion in final good imports, among which $0.88 billion resulted from the decrease in compliance rate during trade war. As for the intermediate goods, the average statutory tariffs for U.S. exports increased to 3.17%. With the total U.S. export of $711 billion to the retaliating countries, the tariff revenue loss for the governments in the retaliating countries was around $13.6 billion in intermediate good imports, among which $5.36 billion resulted from the decrease in compliance during trade war. In summary, the trade-war induced importers to comply with tariffs at a lower rate, generating around $6.2 billion loss of government revenue for U.S. trade partners, which is about 41% of the total loss in tariff revenues due to non-compliance in these countries.

6.3 Product Differentiation

As noted in previous literature, such as (Javorcik and Narciso, 2008), evasion activity may be heightened in trade of differentiated products, relative to homogeneous goods, due to

greater difficulty in assessing unit values. To analyze the potential heterogeneity, we estimate compliance rates separately for differentiated and non-differentiated products. Table 5 and Table 6 show the estimation results. We show a single first stage equation in column (1) of Table 5, which supports the use of the constructed cost shifter as an instrument for prices (F-stat = 78.6). In column (2), the demand elasticities for non-differentiated and differentiated products are similar: around -1.5 for non-differentiated products and around -1.6 for differentiated products.

Table 6 shows the three parameters in the compliance rate functions, as well as the estimated compliance rates in the two periods. The zero-order component, θ_0, is 1.07 for non-differentiated products, and it is only 0.69 for differentiated products, suggesting that compliance at low tariff rates is higher for non-differentiated products. Further, in this case, the compliance function for differentiated products lies below compliance function for non-differentiated products over the entire relevant range. These results highlight the important difference between the two types of the products. Due to the difficulty in determining the product values of differentiated products at importing custom, the tariff compliance rates for differentiated products tend to be lower than that for non-differentiated products when facing the same tariff rates (we discuss reasons for this at the end of this subsection). As in the baseline results, the compliance rate functions indicate a negative marginal effect of increasing tariffs on compliance for both types of products.

Table 6 also show the calculated compliance rates by product differentiation. Compliance for non-differentiated products is higher than for differentiated products in both pre-war periods and post-war periods. Specifically, the pre-war compliance rate for non-differentiated products is 1 and it decreases to 0.85 after the trade war. The pre-war compliance rate for differentiated products is 0.57, and it dropped to 0.33 in the post-war periods. The marginal effect of a tariff increase on compliance is also higher for differentiated products than for non-differentiated products. In the pre-war period, a 1 percentage point increase in tariff rates leads to a 0.9 percentage point decrease in compliance among non-differentiated goods.

Whereas, a 1 percentage point increase in the tariff rate reduces the compliance rate of differentiated products by 1.5 percentage points over the same period. In the post-war periods, a 1 percentage point increase in tariff rates only decrease the compliance rate of non-differentiated products by 0.25 percentage points, while it would reduce the compliance rate of differentiated products by 1 percentage point.

Effective tariff rates and government revenue loss due to non-compliance for differentiated and non-differentiated products are displayed in Table 7.[15] Before the trade war, in 2017, U.S. exports of differentiated products to the retaliating countries totaled $403 billion and the average statutory tariff for U.S. differentiated product exports was 1.81%. Our results suggest the average effective tariff rate was 1.03% when taking non-compliance into consideration. The tariff revenue loss for governments of the retaliating countries was around $3.17 billion from imports of differentiated products. As for the non-differentiated products, the average statutory tariff rate was 2.4% in pre-war periods. There was full compliance prior to the trade war so the effective compliant tariff rate was also 2.4%, and there was no government revenue loss generated from the $662 billions U.S. exports of non-differentiated products to the retaliating countries in 2017.

After the trade war, in 2019, the average statutory tariffs for U.S. exports of differentiated products increased to 2.78%. With total U.S. exports of $423 billion to the retaliating countries, the tariff revenue loss for the governments in the retaliating countries was around $7.84 billion in imports of differentiated products, among which $2.75 billion resulted from decreased compliance during trade war. As for the non-differentiated products, the average statutory tariffs for U.S. exports increased to 5.93%. With total U.S. exports of $129 billion to the retaliating countries, the tariff revenue loss for the governments in the retaliating countries was around $1.15 billion in imports of non-differentiated products, all of which were generated by the decrease in compliance during trade war.

In summary, around 35% of government tariff revenue loss in imports of differentiated

[15]Since we use the product categorization from Rauch (1999), the products in U.S. exports that are unable to categorize in this system are left out of the analysis.

products from U.S. trade partners was generated by the trade war, and the trade war is responsible for 100% of government tariff revenue loss in imports of non-differentiated products. We also note that the average statutory tariff rates for non-differentiated products are much higher than those for differentiated products, which implies government tend to impose higher tariffs on product with higher compliance rates. This policy strategy helped to mitigate the revenue loss from tariff evasion during the trade war.

7 Conclusion

Due to the unobservable nature of tax evasion, developing strategies to identify tax compliance rates is critical to improve our understanding of tax evasion behavior, and to generally inform tax policy, including policies related to tax enforcement. We propose a simple procedure to estimate tax compliance from variation in tax rates. The procedure relies on using market data on prices and quantities, and requires exogenous variation in tax rates and one additional instrument. The procedure is generally applicable to markets that impose taxes on either the supply or demand side. Importantly, it is straightforward to estimate compliance as a function of tax rates, which provides necessary flexibility and lends the model to a variety of settings.

The main obstacle in estimating tax compliance using price and quantity data is separately estimating compliance from the demand or supply elasticity, depending on which side of the market the tax is imposed. The key insight in the proposed method is that compliance rates can be estimated by comparing the effects of exogenous variation in tax rates with the effects of exogenous price variation. More formally, separate identification is achieved by instrumenting for the endogenous prices, which allows for the separate estimation of compliance rates.

Using the variation from tariff increases imposed by US trade partners in 2018, we estimate that compliance rates among the US trade partners was 70 to 80% prior to the tariff

increases. However, we find significant declines in compliance after the additional tariffs were applied, which increased the tariff rates among the treated products in the sample from 7 to 26% in 2018. As a result, tariff compliance in the retaliating countries fell to about 50%. The implied marginal effects of tariff increase on tax evasion suggest that a 1 percentage point increase in tariff rates increases tax evasion by 2 to 10%.

The results highlight the usefulness of estimating tax compliance. Among the retaliatory trade partners in sample, we estimate a combined revenue loss of over $4B in 2017 due to tariff evasion, and an additional $6B in revenue loses after imposing additional tariffs, relative to the pre-war compliance rates. Determining the magnitudes of revenues lost to evasion relies on estimation of the compliance rate levels, which is one major advancement of this work. Further, the sizable loss in tariff revenue is informative about the value of enforcement efforts. Compliance rates and play a key role in determining the costs and benefits of taxation and enforcement.

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8 Tables and Figures

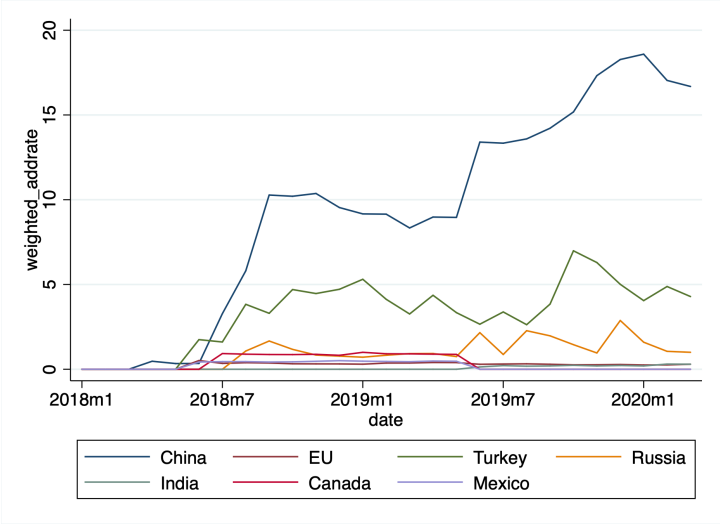
Table 1: The 2018-2019 Trade War

Country	Month enacted	Products	2017 imports		Tariff(%)		
		(#HS-10)	(million $US)	(%)	2017	2018	2019
China	April, 2018	6292	92905	8.5	7.1	10.9	22.6
Mexico	Jun, 2018	115	4500	0.4	0	12.6	9.2
Turkey	Jun, 2018	158	1493	0.1	11.5	19.2	45.8
European Union	June, 2018	188	3909	0.4	3.1	14.5	27.6
Canada	July, 2018	356	18793	1.7	0	6.1	6.0
Russia	August, 2018	117	253	0.0	2.9	11.7	33.0
India	June, 2019	53	622	0.1	18.6	19.1	22.1
Total		7279	122475	11.2	5.7	10.5	19.7

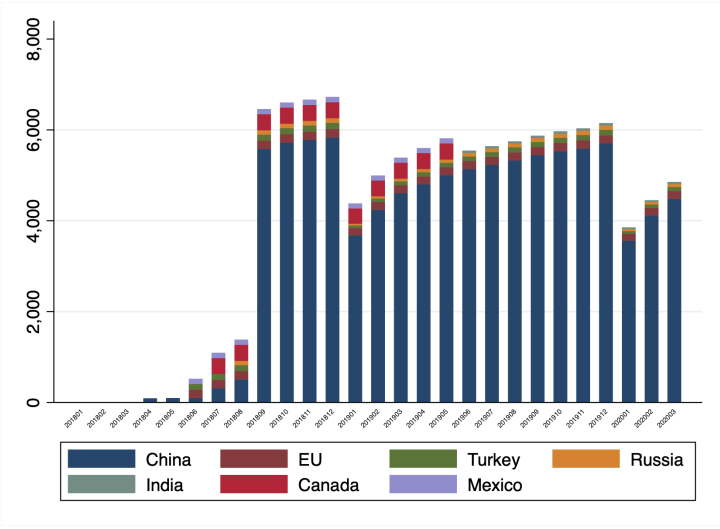
Notes: This table displays country average retaliatory tariff rates. Tariff rates in 2017, 2018 and 2019 by country are computed as annual averages across the retaliating HS-10 products weighted by 2017 trade values. The denominator for import share is the total 2017 annual U.S. exports. See the text for data sources.

Figure 1: Retaliating Tariffs and Coverage

(a) Weighted Average Added Rates

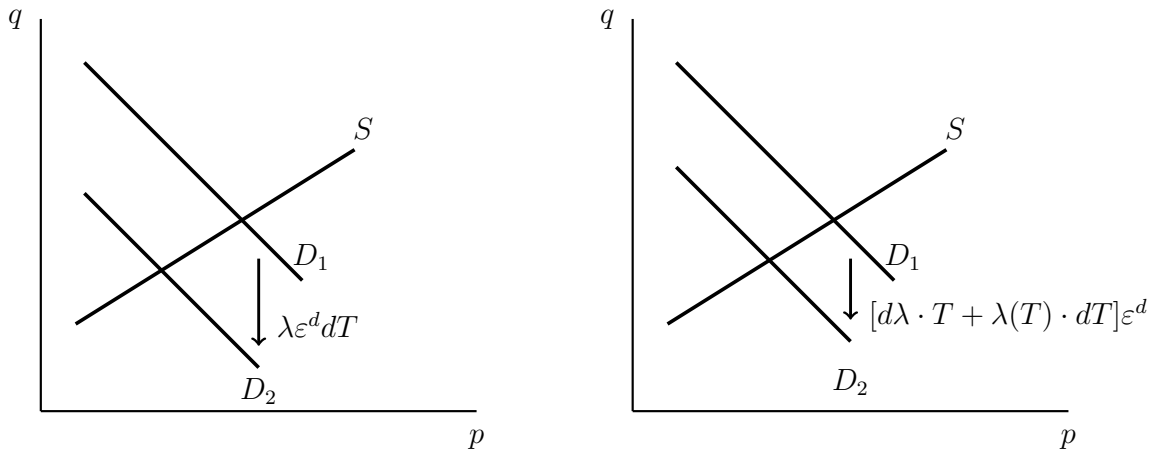


(b) Number of Products in Retaliation



Notes: This figure shows the added tariff rates and the tariff coverage by importing countries from January 2018 to March 2020. Panel (a) shows the line graph of the average added tariff rates, weighted by trade values for each retaliating country. Ordered by the added tariffs in March 2020 from high to low, the retaliating countries include China, Turkey, Russia, India, EU, Canada, and Mexico. Panel (b) shows the bar chart of the numbers of HS-10-level traded products covered under the retaliating tariffs. Each importer-HS-10 pair is counted as a product.

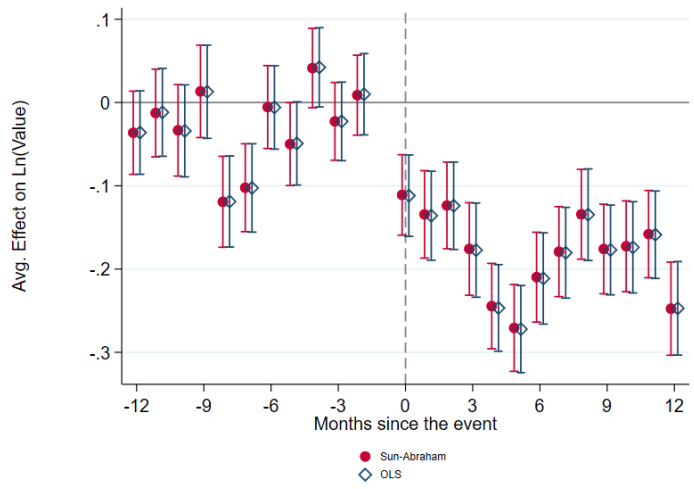
Figure 2: Estimating Compliance with a Tax Increase



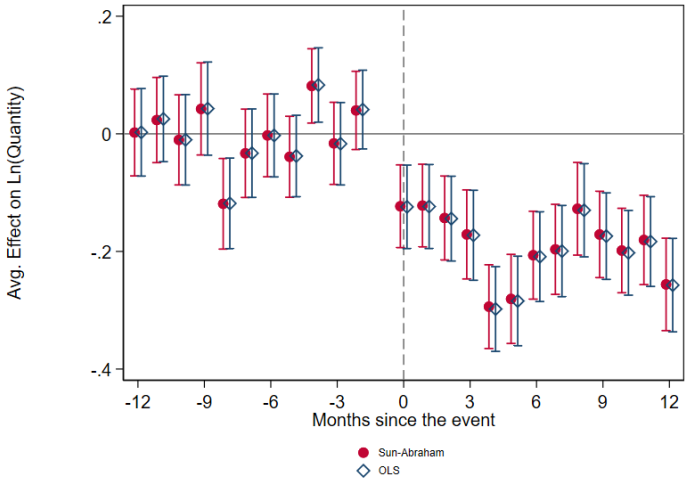
(a) Compliance is Constant in the Tax (b) Compliance Depends on the Tax

Notes: The figure describes the shifts in demand due to increased tax rates when the statutory tax burden is on the demand side under two different scenarios. Prices are tax-exclusive; demand shifts reflect changes in tax-inclusive prices faced by importers. In Panel (a), the compliance rate is constant, i.e., it does not depend on the tax rate. In that case, the magnitude of the demand shift depends on the elasticity of demand and the constant rate of compliance. Panel (b) includes a more general case, in which the compliance rate may depend on the tax rate. In Panel (b), the magnitude of the shift includes a potential change in compliance, which mitigates the magnitude of the shift from a tax increase, assuming that compliance is negatively related to the tax rate.

Figure 3: Event Studies, Binary Treatment



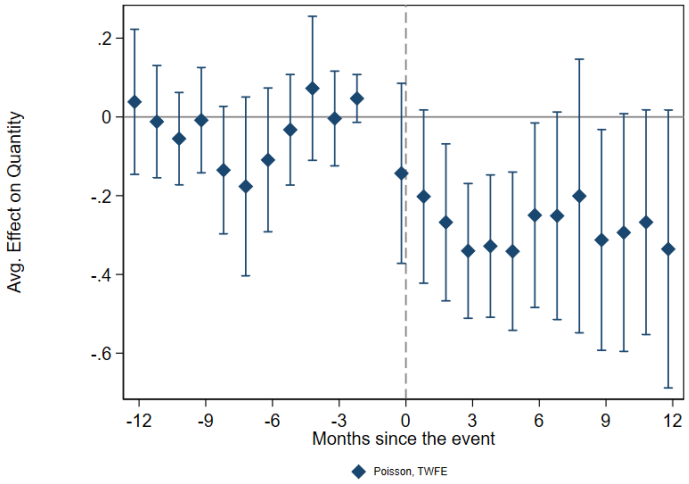
(a) Effects of Added Tariff on ln(Value).



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

Event study estimated effects for having any added rate on ln(value) and ln(quantity). Observations are at the Country by HS-10 by month level. Each panels includes two sets of results: One from an OLS estimator with two-way fixed effects, and another with the Sun and Abraham analog. Each includes Country-HS-10 and month fixed effects. Dropped the highest value HS10 product in each HS6 category for treated HS6 groups. Confidence intervals based on clustering at the Country-HS6 level.

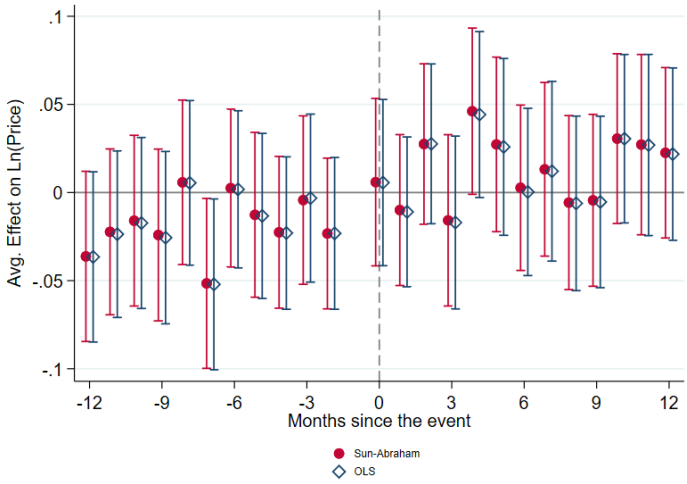
Figure 4: Poisson Event Studies, Binary Treatment



(a) Effect of Added Tariff on Quantity, Poisson

Poisson event study estimated effects for having any added rate on quantities. Observations are at the Country by HS-10 by month level. Poisson estimator with two-way fixed effects. Each includes Country-HS-10 and month fixed effects. Dropped the highest value HS10 code within HS6 in treated HS6 codes. Confidence intervals based on clustering at the Country-HS6 level.

Figure 5: Event Studies, Binary Treatment



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

Event study estimated effects for having any added rate on ln(price). Observations are at the Country by HS-10 by month level. Each panels includes two sets of results: One from an OLS estimator with two-way fixed effects, and another with the Sun and Abraham analog. Each includes Country-HS-10 and month fixed effects. No weights are included. Dropped highest value HS10 products within treated HS6 categories. Confidence intervals based on clustering at the Country-HS6 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	0.258	−1.677***	0.076	−1.468***
	(0.231)	(0.404)	(0.188)	(0.291)
T^2	−0.719	4.132**	−0.083	3.567**
	(0.970)	(1.758)	(0.767)	(1.190)
T^3	0.332	−3.680*	−0.080	−2.990**
	(0.986)	(1.930)	(0.800)	(1.219)
Share × $ln(Price^f)$	4.947***		4.018***	
	(0.685)		(0.454)	
ln(Price)		−1.736***		−1.729***
		(0.172)		(0.141)
Mean	4.154	7.148	4.109	6.756
F-Stat	52.218		78.218	
Observations	1521158	1521158	3461091	3461091
Clusters	55677	55677	159520	159520
$\hat{\theta}_0$		0.966		0.849
$SE(\hat{\theta}_0)$		0.264		0.186
$\hat{\theta}_1$		−2.38		−2.063
$SE(\hat{\theta}_1)$		1.063		0.715
$\hat{\theta}_2$		2.119		1.729
$SE(\hat{\theta}_2)$		1.143		0.72
$\hat{\lambda}_1$		0.809		0.712
$SE(\hat{\lambda}_1)$		0.2		0.144
$\hat{\lambda}_2$		0.522		0.458
$SE(\hat{\lambda}_2)$		0.1		0.078
Country-HS-10 FE	x	x	x	x
Country-Month-Year FE	x	x	x	x

Dropped highest value HS10 within each HS6 groups. Pooled post periods. The table reports the reduced-form estimates of the effect of tariff rates (T_{jht}) and the cost shifter on trade quantities and prices, as well as the corresponding structural estimates. The only additional controls are the fixed effects indicated near the bottom of the table: Country by HS-10 and country-month-year fixed effects. Columns (1) and (2) include varieties imported by retaliating countries only. Columns (3) and (4) include all varieties exported by the U.S. The corresponding estimates of the structural parameters are calculated as described and included in the bottom panel. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 3: Reduced Form, Intermediate and Final Goods

	(1) *Pooled First Stage*	(2) IV
	ln(Price)	ln(Quantity)
1[Final Good]		
$\times T$	0.078	−3.085**
	(0.441)	(1.322)
$\times T^2$	−0.871	8.637*
	(1.751)	(5.244)
$\times T^3$	0.936	−7.188
	(1.953)	(6.374)
1[Intermediate Good]		
$\times T$	0.010	−1.281***
	(0.215)	(0.325)
$\times T^2$	0.504	3.456**
	(0.928)	(1.401)
$\times T^3$	−0.630	−3.075**
	(1.003)	(1.391)
Share $\times$ $ln(price^f)$	3.999***	
	(0.454)	
$ln(price)$		
$\times$ 1[Final Good]		−3.331***
		(0.734)
$\times$ 1[Intermediate Good]		−1.695***
		(0.161)
Mean	4.109	6.757
F-Stat	77.423	
Observations	3459471	3459471
Clusters	159481	159481
Country-HS-10 FE	x	x
Country-Month-Year FE	x	x

Dropped highest value HS10 within each HS6 groups. The table reports the reduced-form estimates for products that considered intermediate (more likely to be eligible to bypass tariffs) and final goods (less likely to be eligible to bypass tariffs) by using free trade zones. Column (1) includes first stage estimates using a single common shifter. Column (2) reports the second stage estimates while allowing for different demand elasticities. In this case, we use Z and ln(price) interacted with heterogeneity categories. The only additional controls are the fixed effects indicated near the bottom of the table: Country by HS-10 and country-month-year fixed effects. The corresponding estimates of the structural parameters are included in the following table. The structural estimates are calculated as described. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 4: Compliance, Intermediate and Final Goods

	(1)	(2)
	$1[FinalGood]$	$1[IntGood]$
$\hat{\theta}_0$	0.926	0.756
$SE(\hat{\theta}_0)$	0.46	0.207
$\hat{\theta}_1$	−2.593	−2.039
$SE(\hat{\theta}_1)$	1.706	0.842
$\hat{\theta}_2$	2.158	1.814
$SE(\hat{\theta}_2)$	1.995	0.829
$\hat{\lambda}_1$	0.7	0.634
$SE(\hat{\lambda}_1)$	0.327	0.162
$\hat{\lambda}_2$	0.366	0.396
$SE(\hat{\lambda}_2)$	0.158	0.09

The table reports the reduced-form estimates for products that considered intermediate (more likely to be eligible to bypass tariffs) and final goods (less likely to be eligible to bypass tariffs) by using free trade zones. Estimates based on the results in Column (2) (IV) from the previous table, allowing for different demand elasticities for final and intermediate goods. We use Z interacted with heterogeneity categories to instrument for ln(price) interacted with heterogeneity categories. The only additional controls are the fixed effects indicated near the bottom of the table: Country by HS-10 and country-month-year fixed effects. The corresponding estimates of the reduced form parameters are included in the previous table. The structural estimates are calculated as described. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 5: Reduced Form, By Product Differentiation

	(1) *Pooled First Stage*	(2) IV, Different ε_d
	ln(Price)	ln(Quantity)
1[Non-Differentiated]		
$\times T$	0.049	−1.597**
	(0.242)	(0.487)
$\times T^2$	−0.503	1.819
	(0.996)	(2.111)
$\times T^3$	0.213	−1.815
	(1.022)	(2.149)
1[Differentiated]		
$\times T$	0.242	−1.129**
	(0.273)	(0.364)
$\times T^2$	−0.508	2.954**
	(1.014)	(1.306)
$\times T^3$	0.336	−1.902*
	(0.933)	(1.061)
Share $\times$ $ln(Price^f)$	4.031***	
	(0.454)	
$ln(price)$		
$\times$ 1[Non-Differentiated]		−1.488***
		(0.154)
$\times$ 1[Differentiated]		−1.641***
		(0.214)
Mean	4.109	6.756
F-Stat	78.662	
Observations	3461091	3461091
Clusters	159520	159520
Country-HS-10 FE	x	x
Country-Month-Year FE	x	x

The table reports the reduced-form estimates for non-differentiated and differentiated products. Column (1) includes first stage estimates using a single common shifter (Estimating a single ε_d) and Column (2) displays the corresponding second stage estimates allowing for different demand elasticities. In this case, we use Z and ln(price) interacted with heterogeneity categories. The only additional controls are the fixed effects indicated near the bottom of the table: Country by HS-10 and country-month-year fixed effects. The corresponding estimates of the structural parameters are included in the following table. The structural estimates are calculated as described. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6: Compliance, By Product Differentiation

	(1) Non-Differentiated	(2) Differentiated
$\widehat{\theta_0}$	1.073	0.688
$SE(\hat{\theta}_0)$	0.349	0.255
$\widehat{\theta_1}$	−1.222	−1.8
$SE(\widehat{\theta_1})$	1.428	0.855
$\widehat{\theta_2}$	1.219	1.159
$SE(\widehat{\theta_2})$	1.451	0.676
$\widehat{\lambda}_1$	1.006	0.567
$SE(\widehat{\lambda}_1)$	0.28	0.203
$\widehat{\lambda}_2$	0.849	0.333
$SE(\widehat{\lambda}_2)$	0.148	0.116

The table reports the structural estimates for non-differentiated and differentiated products. Columns (1) and (2) include estimates using a single common shifter (Estimating a single ε_d) and Columns (3) and (4) display estimates while allowing for different demand elasticities. In this case, we use Z and ln(price) interacted with heterogeneity categories. The only additional controls are the fixed effects indicated near the bottom of the table: Country by HS-10 and country-month-year fixed effects. The corresponding estimates of the reduced form parameters are included in the previous table. The structural estimates are calculated as described. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 7: Effective Tariffs by Product Types

Panel (a) Compliance and Effective Tariffs in 2017

	(1) Statutory Tariffs (%)	(2) Compliance Rate	(3) Effective Tariffs (%)	(4) U.S. Export Values($Billion)	(5) Tariff Revenue Loss ($Billion)
Final Goods	1.355	0.700	0.949	76.7	0.31
Intermediate Goods	1.833	0.634	1.162	662.0	4.44
Differentiated Products	1.814	0.567	1.029	403.0	3.17
Non-differentiated Products	2.396	1.000	2.396	126.0	0

Panel (b) Compliance and Effective Tariffs in 2019

	(1) Statutory Tariffs (%)	(2) Compliance Rate	(3) Effective Tariffs (%)	(4) U.S. Export Values($Billion)	(5) Tariff Revenue Loss ($Billion)	(6) War-induced Revenue Loss ($Billion)
Final Goods	3.269	0.366	1.196	80.3	1.66	0.88
Intermediate Goods	3.165	0.396	1.253	711.0	13.59	5.36
Differentiated Products	2.779	0.333	0.925	423.0	7.84	2.75
Non-differentiated Products	5.927	0.849	5.032	129.0	1.15	1.15

This table shows the export-value-weighted average compliance rate and effective tariffs in 2017 and 2019 for different product types. The compliance rates are from Table 4 and 6. The effective tariffs is the product of statutory tariffs and the estimated compliance rates. Given U.S. export values, the tariff revenue loss is calculated from the average statutory tariffs and compliance rates. The war-induced revenue loss is calculated using the change in compliance rate during 2017-2019 and the 2019 average statutory tariffs, given U.S. export values in 2019.

9 Appendix

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:

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

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):

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

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:

$$\Pi = P(Q)q - cq^{\phi},$$

with $\phi > 1$ to ensure that the second-order conditions of profit maximization hold.

We use the conduct parameter approach, where $0 < \theta < 1$ refers to conduct. This nests a Cournot model with $\theta = 1/N$ so that $\theta Q = q$ in equilibrium. The first-order condition for

profit maximization is:

$$\begin{aligned} P - c\phi q^{\phi-1} + \frac{P}{Q\varepsilon^d} q &= 0, \\ P - c\phi(\theta Q)^{\phi-1} + \frac{P\theta}{\varepsilon^d} &= 0, \\ Q = \frac{1}{\theta}\left[\frac{1}{c\phi}\left(1 + \frac{\theta}{\varepsilon^d}\right)P\right]^{\frac{1}{\phi-1}}. \end{aligned}$$

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:

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

Using logs and simplifying, we obtain the following system that mirrors the structure of the competitive case:

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

The conduct parameter θ and cost parameter c are absorbed into the constant term of the supply-equivalent equation and the error terms, leaving the slope coefficients, which identify our parameters of interest, unaffected by market structure. 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:

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

The reduced form coefficients are:

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

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

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

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

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

Table A1: Out-of-Sample Model Selection

Polynomial Order (K)	0 (Linear)	1 (Quadratic)	2 (Cubic)	3 (Quartic)
OOS MSE	253.32	252.88	**252.36**	252.45

Appendix B: Supplemental Tables and Figures

Table B.1: Tax-Dependent Compliance Rate Estimates - Alternate Fossil Shifters

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Contemporaneous, Share $\times$ $ln(Price_t^f)$				*Contemporaneous, Share* $\times$ $(Price_t^f/Price_{2016}^f)$			
	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	0.260	-1.730***	0.076	-1.484***	0.260	-1.730***	0.077	-1.485***
	(0.231)	(0.390)	(0.188)	(0.280)	(0.231)	(0.390)	(0.188)	(0.279)
T^2	-0.725	4.284**	-0.085	3.592**	-0.726	4.285**	-0.086	3.595**
	(0.970)	(1.699)	(0.767)	(1.139)	(0.970)	(1.698)	(0.767)	(1.135)
T^3	0.338	-3.758**	-0.077	-2.988**	0.339	-3.759**	-0.077	-2.987**
	(0.986)	(1.857)	(0.800)	(1.153)	(0.986)	(1.856)	(0.799)	(1.147)
Share $\times$ $ln(Price_t^f)$	5.357***		4.432***					
	(0.698)		(0.463)					
Share $\times$ $(Price_t^f/Price_0^f)$					3.847***		3.196***	
					(0.509)		(0.338)	
ln(Price)		-1.555***		-1.576***		-1.554***		-1.561***
		(0.148)		(0.121)		(0.150)		(0.122)
Mean	4.154	7.148	4.109	6.756	4.154	7.148	4.109	6.756
F-Stat	58.989		91.558		57.223		89.558	
Observations	1521158	1521158	3461091	3461091	1521158	1521158	3461091	3461091
Clusters	55677	55677	159520	159520	55677	55677	159520	159520
$\hat{\theta}_0$		1.112		.941		1.113		.951
$SE(\hat{\theta}_0)$		.285		.196		.286		.198
$\hat{\theta}_1$		-2.754		-2.278		-2.757		-2.303
$SE(\hat{\theta}_1)$		1.148		.75		1.151		.755
$\hat{\theta}_2$		2.417		1.895		2.419		1.913
$SE(\hat{\theta}_2)$		1.229		.748		1.231		.752
$\hat{\lambda}_1$		.93		.79		.931		.798
$SE(\hat{\lambda}_1)$		.216		.152		.217		.153
$\hat{\lambda}_2$		.597		.509		.598		.514
$SE(\hat{\lambda}_2)$		.108		.082		.108		.083
Country-HS-10 FE	x	x	x	x	x	x	x	x
Country-Month-Year FE	x	x	x	x	x	x	x	x

Dropped highest value HS10 within each HS6 groups. Pooled post periods. The table reports the reduced-form and structural estimates from using contemporaneous fossil cost shifters. The first 4 columns use the cost share from 2016 interacted with the ln(fossil price) at time t, and the last 4 use the cost share in 2016 interacted with price at time t relative to 2016. Fixed effects are indicated at the bottom of the table. All include Country-HS10 and month-year fixed effects. Columns (1) and (2) and (5)-(6) include varieties imported by retaliating countries only. Columns (3) and (4) and (7)-(8) include all varieties exported by the U.S. The corresponding estimates of the structural parameters are calculated as described and included in the bottom panel. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.2: 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	0.384*	-1.743***	0.215	-1.407***	0.148	-1.687***	-0.015	-1.521***
	(0.222)	(0.379)	(0.179)	(0.271)	(0.230)	(0.416)	(0.189)	(0.303)
T^2	-0.911	4.478**	-0.326	3.281**	-0.352	4.182**	0.265	3.742**
	(0.895)	(1.621)	(0.703)	(1.093)	(0.965)	(1.817)	(0.780)	(1.254)
T^3	0.489	-4.051**	0.096	-2.804**	0.044	-3.743*	-0.345	-3.132**
	(0.906)	(1.803)	(0.721)	(1.121)	(0.977)	(2.009)	(0.826)	(1.306)
Share $\times$ $ln(Price^f)$	4.724***		3.829***		4.224***		3.463***	
	(0.677)		(0.450)		(0.929)		(0.618)	
ln(Price)		-1.711***		-1.686***		-1.861***		-1.867***
		(0.177)		(0.144)		(0.298)		(0.239)
Mean	4.154	7.148	4.109	6.756	4.154	7.148	4.109	6.756
F-Stat	48.74		72.406		20.67		31.405	
Observations	1521158	1521158	3461211	3461211	1521158	1521158	3461091	3461091
Clusters	55677	55677	159539	159539	55677	55677	159520	159520
$\hat{\theta}_0$		1.018		.834		.906		.814
$SE(\hat{\theta}_0)$		.264		.185		.28		.192
$\hat{\theta}_1$		-2.616		-1.945		-2.246		-2.004
$SE(\hat{\theta}_1)$		1.018		.68		1.064		.708
$\hat{\theta}_2$		2.367		1.662		2.01		1.677
$SE(\hat{\theta}_2)$		1.099		.683		1.132		.72
$\hat{\lambda}_1$		.846		.705		.758		.681
$SE(\hat{\lambda}_1)$		.202		.144		.217		.151
$\hat{\lambda}_2$		.533		.467		.488		.435
$SE(\hat{\lambda}_2)$		.103		.079		.116		.087
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

Dropped highest value HS10 within each HS6 groups. Pooled post periods. The table reports the reduced-form and structural estimates from using different fixed effects specifications. Fixed effects are indicated at the bottom of the table. Columns (1)-(4) include Country-HS10 and month-year fixed effects. Columns (5)-(8) include Country-HS10, Country-Month, and Industry-Month fixed effects. Columns (1) and (2) and (5)-(6) include varieties imported by retaliating countries only. Columns (3) and (4) and (7)-(8) include all varieties exported by the U.S. The corresponding estimates of the structural parameters are calculated as described and included in the bottom panel. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.3: 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]	0.048*			0.048*		
	(0.025)			(0.025)		
Max Added Rate in HS4		0.289**			0.289**	
		(0.123)			(0.123)	
Mean Added Rate in HS4			0.550*			0.550*
			(0.296)			(0.296)
Observations	1441832	1441832	1441832	3381765	3381765	3381765
Clusters	53847	53847	53847	157690	157690	157690
Country-HS10 FE	x	x	x	x	x	x
Country-Month-Year FE	x	x	x	x	x	x

Estimated effects of added rates within HS4-country group on the untreated products. Each sample includes untreated HS10s only. All treatments vary at the HS4-month level. Dropped highest value HS10 within each HS6 groups. The only additional controls are the fixed effects indicated near the bottom of the table: Country by HS-6 and country-month-year fixed effects. Columns (1)-(3) include varieties imported by retaliating countries only. Columns (4)-(6) include all varieties exported by the U.S. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.4: Estimating λ with Misclassification

	(1)	(2)	(3)	(4)
	Retaliating Countries Only		*All Sample Countries*	
	ln(Price)	ln(Quantity)	ln(Price)	ln(Quantity)
T	0.222	-1.425***	0.104	-1.194***
	(0.194)	(0.338)	(0.150)	(0.236)
T^2	-0.551	2.426**	-0.236	1.941**
	(0.600)	(1.127)	(0.406)	(0.712)
T^3	0.155	-2.226	0.112	-1.322**
	(0.696)	(1.456)	(0.408)	(0.658)
ln(1+Mean Added Rate in HS4)	0.035	-0.010	0.030	-0.074
	(0.093)	(0.147)	(0.091)	(0.144)
Share $\times$ $ln(Price^f)$	4.950***		4.018***	
	(0.685)		(0.454)	
$ln(Price)$		-1.738***		-1.731***
		(0.172)		(0.141)
Mean	4.154	7.148	4.109	6.756
F-Stat	52.282		78.225	
Observations	1521158	1521158	3461091	3461091
Clusters	55677	55677	159520	159520
$\hat{\theta}_0$		.82		.689
$SE(\hat{\theta}_0)$		.221		.152
$\hat{\theta}_1$		-1.396		-1.121
$SE(\hat{\theta}_1)$		.679		.427
$\hat{\theta}_2$		1.281		.763
$SE(\hat{\theta}_2)$		.853		.387
$\hat{\lambda}_1$		.728		.614
$SE(\hat{\lambda}_1)$		.181		.127
$\hat{\lambda}_2$		.562		.467
$SE(\hat{\lambda}_2)$		.116		.083
Country-HS10 FE	x	x	x	x
Country-Month-Year FE	x	x	x	x

The table reports the reduced-form estimates of the effect of tariff rates (T_{jht}) and the cost shifter on trade quantities and prices, as well as the corresponding structural estimates, conditional on the mean added rate in the HS4, excluding products in own HS6. Calculated with fixed sample of products that includes all HS10s-Country products in data over entire sample period. The only additional controls are the fixed effects indicated near the bottom of the table: Country by HS-10 and country-month-year fixed effects. Columns (1) amd (2) include varieties imported by retaliating countries only. Columns (3) and (4) include all varieties exported by the U.S. The corresponding estimates of the structural parameters are calculated as described and included in the bottom panel. Dropped highest value HS10 within each HS6 groups. Pooled post periods. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Country Heterogeneity

Table B.5: Country Heterogeneity

	China		*EU + Canada*		*CH, MX, RU, TR*	
	ln(price)	ln(Quantity)	ln(price)	ln(Quantity)	ln(price)	ln(Quantity)
τ	-0.074	-0.900**	0.001	-1.296***	-0.124	-0.952***
	(0.228)	(0.359)	(0.202)	(0.302)	(0.186)	(0.284)
τ^2	0.024	1.590*	0.372	2.348**	0.039	1.596**
	(0.485)	(0.904)	(0.565)	(0.902)	(0.434)	(0.753)
τ^3	0.239	-1.128**	0.056	-1.074**	0.102	-1.214*
	(0.315)	(0.532)	(0.379)	(0.424)	(0.396)	(0.623)
Share × ln(Price)	3.360***		3.839***		3.710***	
	(0.594)		(0.492)		(0.544)	
lnprice		-1.710***		-1.879***		-1.564***
		(0.222)		(0.169)		(0.177)
Mean	4.075	6.496	4.158	6.584	4.021	6.713
F-Stat	32.023		60.867		46.436	
Observations	1997737	1997737	3051695	3051695	2269099	2269099
Clusters	105168	105168	148052	148052	112705	112705
$\hat{A}$		.526		.689		.608
$SE(\hat{A})$		.219		.174		.19
$\hat{B}$		-.929		-1.249		-1.02
$SE(\hat{B})$		.544		.489		.497
$\hat{C}$		.659		.571		.776
$SE(\hat{C})$		.332		.233		.412
$\hat{\lambda}_1$		.454		.631		.533
$SE(\hat{\lambda}_1)$		.182		.153		.158
$\hat{\lambda}_2$		.356		.402		.419
$SE(\hat{\lambda}_2)$		.135		.085		.113
Country-HS-10 FE	x	x	x	x	x	x
Country-Month-Year FE	x	x	x	x	x	x

IV and Structural Estimates for subgroups of countries. Estimates from separate regressions. Each sample includes the treated countries referenced in the column heading plus all control countries. Dropped highest value HS10 within each HS6 groups. Pooled post periods. The top panel reports the reduced-form estimates of the effect of tariff rates (T_{jht}) and the cost shifter on trade quantities and prices. The bottom panel includes the corresponding structural estimates. The only additional controls are the fixed effects indicated near the bottom of the table: Country by HS-10 and country-month-year fixed effects. The corresponding estimates of the structural parameters are calculated as described and included in the bottom panel. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.6: Country Heterogeneity by Final/Intermediate Goods

	China	*EU + Canada*	*CH, MX, RU, TR*
	ln(Quantity)	ln(Quantity)	ln(Quantity)
$ln(price) \times 1[intgood]$	-3.759**	-3.488***	-3.559**
	(1.518)	(0.791)	(1.476)
$ln(price) \times 1[finalgood]$	-1.504***	-1.840***	-1.399***
	(0.267)	(0.191)	(0.216)
$\tau \times 1[intgood]$	-0.802	-4.629**	-0.629
	(1.901)	(1.700)	(1.562)
$\tau^2 \times 1[intgood]$	1.422	15.698**	-1.781
	(7.072)	(6.610)	(6.431)
$\tau^3 \times 1[Intgood]$	-3.105	-13.786**	3.737
	(7.469)	(6.513)	(7.365)
$\tau \times 1[finalgood]$	-1.326**	-1.086**	-1.330***
	(0.450)	(0.363)	(0.384)
$\tau^2 \times 1[finalgood]$	3.867**	2.754*	3.660**
	(1.506)	(1.456)	(1.428)
$\tau^3 \times 1[finalgood]$	-3.361**	-2.018**	-3.216**
	(1.362)	(0.992)	(1.331)
Mean	6.497	6.584	6.714
Observations	1996799	3050757	2267479
Clusters	105134	148018	112666
Country-HS-10 FE	x	x	x
Country-Month-Year FE	x	x	x

	China		*EU + Canada*		*CH, MX, RU, TR*	
	1[finalgood]	1[intgood]	1[finalgood]	1[intgood]	1[finalgood]	1[intgood]
$\hat{A}$	.213	.881	1.327	.59	.176	.95
$SE(\hat{A})$	.537	.327	.528	.207	.463	.301
$\hat{B}$	-.378	-2.57	-4.499	-1.497	.5	-2.615
$SE(\hat{B})$	1.945	1.074	1.924	.792	1.712	1.072
$\hat{C}$	.825	2.234	3.952	1.097	-1.05	2.298
$SE(\hat{C})$	2.121	.999	1.878	.543	1.856	.995
$\hat{\lambda}_1$	.179	.699	1.006	.56	.221	.773
$SE(\hat{\lambda}_1)$	.346	.261	.401	.193	.305	.237
$\hat{\lambda}_2$	.171	.445	.366	.296	.236	.496
$SE(\hat{\lambda}_2)$	.206	.177	.173	.092	.209	.15

IV Estimates for subgroups of countries by product type, intermediate vs. final goods. Estimates from separate regressions. Each sample includes the treated countries referenced in the column heading plus all control countries. Dropped highest value HS10 within each HS6 groups. Pooled post periods. The top panel reports the reduced-form estimates of the effect of tariff rates (T_{jht}) and the cost shifter on trade quantities and prices. The bottom panel includes the corresponding structural estimates. The only additional controls are the fixed effects indicated near the bottom of the table: Country by HS-10 and country-month-year fixed effects. The corresponding estimates of the structural parameters are calculated as described and included in the bottom panel. Standard errors clustered at the Country-HS6 level are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.