Do Foreigners Pay for Tariffs?*

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Abstract

Tariffs are back, and with them the century-old question about who pays for them. In this paper, we provide direct evidence about whether foreigners do so, as well as about the magnitude of the associated transfers. We implement a difference-in-difference strategy to estimate how cross-country differences in exposure to global tariff changes affect countries' relative export prices. Crucially, this method compares the prices of what countries sell to and buy from the rest of the world. Using data on the near universe of worldwide export prices and tariff changes between 1995 and 2021, we find large and statistically significant declines in the terms of trade of countries exposed to tariffs imposed by their trading partners, with most of the adjustment accounted for by exchange rate fluctuations.

*This version: March 2026. Author contacts: rodrigo.adao@chicagobooth.edu, keelanb@mit.edu, costinot@mit.edu, and ddonald@mit.edu. Preliminary and incomplete. Please do not circulate.

1 Introduction

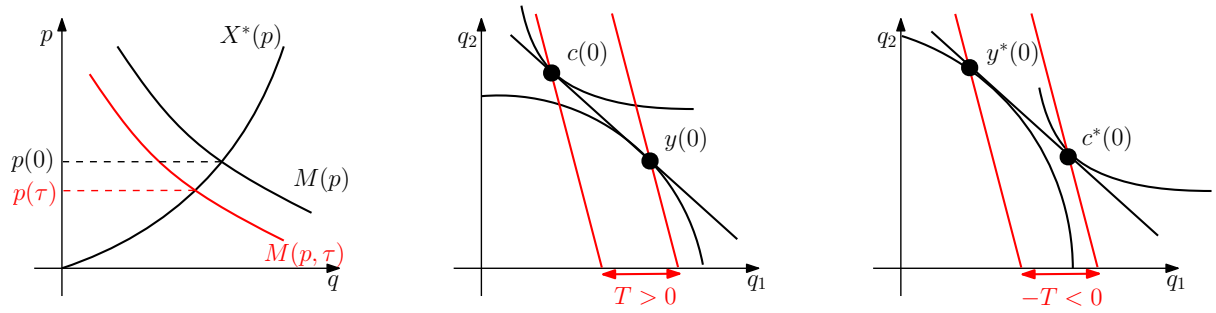
Tariffs are back, and with them the century-old question about who pays for them. As a matter of theory, the mechanism through which foreigners may share some of the burden of another country's tariff is well understood. Mill (1844), Torrens (1844), and Edgeworth (1894) offer early discussions. By imposing a tariff, a country may lower its import demand and, in turn, induce a decline in the price of what it buys relative to what it sells. This terms-of-trade adjustment creates an income transfer from foreigners to the tariff-imposing country, as illustrated in Figure 1. The goal of this paper is to provide empirical evidence on whether such terms-of-trade movements occur in practice and, if so, on how large the associated transfers are.

Two main lines of research have been previously used to shed light on this question. Both are informative, but none of them is fully satisfactory. The first one is structural and consists of writing a complete model of the world economy in order to, either through estimation or calibration, obtain knowledge of the counterparts of the import demand and export supply curves in Figure 1. Given estimates of these curves, counterfactual analysis can then shed light on the full incidence of tariffs on foreigners. Examples of this approach include Broda et al. (2008), Costinot and Rodríguez-Clare (2014), Ossa (2014), and Caliendo and Parro (2015). These papers suggest large terms-of-trade effects from tariffs, with the caveat that no actual changes in tariffs enter the empirical analysis.

The second approach is reduced-form and focuses on the differential effect of tariffs on import prices across products or source countries. Amiti et al. (2019) and Fajgelbaum et al. (2020) are two recent and influential examples. In the context of the first US-China trade war, they compare the prices received by foreign exporters that were directly targeted by US tariffs to those that were not. They find no statistically significant difference in the evolution of the two sets of prices. This striking fact is consistent with foreigners not paying for US tariffs. But since the empirical analysis only compares the relative price of various US imports, not the relative price of exports to imports described in Figure 1, this is equally consistent with foreigners paying for US tariffs to a significant extent.[1]

In this paper, we propose an alternative reduced-form approach that maps directly to the causal mechanism highlighted in Figure 1. Our empirical design is a difference-in-difference specification, implemented across countries and over time. At each horizon, it asks how cross-country differences in exposure to global tariff changes affect countries' average export prices relative to those of their trading partners. Crucially, this relative

[1] In fact, when combining their reduced-form evidence with a structural model, Fajgelbaum et al. (2020) conclude that, absent foreign retaliation, US tariffs would have lead to a non-trivial terms-of-trade improvement and a positive transfer from the rest of the world.



(a) Price Declines (b) Home Gains (c) Foreign Loses

Figure 1: When Foreigners Pay for the Tariff

Notes: Panel 1a describes how a tariff τ affects the price $p(\tau)$ received by foreigners by shifting down domestic import demand $M(\cdot, \tau)$, given foreign export supply $X^*(\cdot)$. Panel 1b describes the associated transfer T for the tariff-imposing country, with its trade balance condition before the tariff in black and after the tariff in red. Panel 1c describes the opposite transfer in the rest of the world.

price compares the prices of what a country sells to and buys from the rest of the world. We use it as a proxy for its terms of trade. Our preferred estimate of the causal effect of global tariff changes on a country's terms of trade is given by the impact of its own tariff exposure minus the impact of its trading partners' exposure, weighted by their shares in its import basket.

We implement this approach on what is essentially the universe of data on product-level bilateral export prices and tariffs around the world, from 1995 through 2021. In order to isolate the causal effects of plausibly exogenous tariff changes we deploy an identification strategy that builds on the approach of Boehm et al. (2023) in their estimation of how trade volumes respond to tariffs. As pointed out by those authors, the WTO's most-favored nation (MFN) principle implies that, whenever two member countries have not negotiated a preferential trade agreement (PTA), the tariff that the importer charges on the exporter must reflect the MFN constraint: it must be the same for all other (non-PTA) exporters. To the extent that MFN tariff changes are endogenously driven by economic conditions in the exporters' markets, we presume that these are driven by the relatively large exporters' conditions since these matter more for the importer that is deciding on its tariff. Yet the MFN constraint means that smaller exporters are also subject to any such tariff changes, so they experience tariff changes as if by accident.[2] Our approach

[2]This is analogous to what Redding and Turner (2015) refer to as the "inconsequential place approach" to research designs aiming to estimate the impact of transportation infrastructure investments. For example, roads are constrained to follow relatively straight lines, so whenever they are constructed in order to connect potentially endogenously selected cities they also end up quasi-accidentally treating locations that lie near the straight line between those cities.

defines thresholds for "small" (based on simultaneous conditions based on absolute size and size within product groups) and then estimates the effect on exporters' export prices of changes in their exposure to MFN tariffs among destinations and products for which the exporter is "small".

Our main empirical findings are as follows. First, we estimate that countries exposed to global tariff changes experience significant downward adjustments in their terms of trade. In fact, these effects are roughly one-for-one within about three years, meaning that if an exporter were to face a 1% increase in the tariff it faces in all destinations and products then its terms of trade would deteriorate by approximately 1%.

This finding from a macro-level approach is very different from the small price effects of tariffs estimated in an emerging body of micro-level studies. Yet, as discussed above, these findings are not contradictory. Micro-level studies have estimated effects of tariffs on certain relative prices, such as the price an importer faces from one exporter relative to another exporter. And those relative effects could be small even if the absolute effect of the importer's tariffs on the price of its imports relative to its exports is large. Indeed, consistent with this dichotomy we replicate, in our worldwide sample, the same types of small relative effects that have been found in country-specific micro-level studies.

Our second main finding is that the large aggregate effect of tariffs on exporters' prices that we estimate takes place largely through exchange rate adjustments. That is, the path of estimated effects on export prices in US dollars is similar to the corresponding path on the exporter's exchange rate relative to the dollar. Put differently, we find essentially no effect of tariffs on export prices when measured in the local currency of the exporter.

The rest of the paper is organized as follows. Section 2 provides the conceptual framework for our main regression. Section 3 describes the empirical strategy that we deploy and the data we use to implement it. Finally, Section 4 presents our empirical estimates.

2 Conceptual Framework

In this section, we start by describing how the mechanism in Figure 1 extends to a setting with many goods and many countries. We then propose a reduced-form specification that maps to this mechanism. Throughout, we emphasize price adjustments at the macro-level, which is the dimension that existing reduced-form approaches abstract from.

2.1 Economic Environment

Consider a world economy with many countries, indexed by o and d, and many products, indexed by g. Let p_{odg} denote the price of product g received by producers from country o selling in country d. Each destination country d imposes ad-valorem tariffs $\tau_d \equiv \{\tau_{odg}\}$ that can vary across origins and products. Let $\tau \equiv \{\tau_d\}$ denote the vector of global tariffs.

To isolate the role of price adjustments at the macro-level, we assume that the supply-side of the economy is Ricardian. In each country, goods are produced using a single composite factor under constant returns to scale. $\bar{L}_o$ denotes the fixed supply of this factor in country o and w_o denotes its price. All markets are perfectly competitive.

The equilibrium conditions that determine prices, $w \equiv \{w_o\}$ and $p \equiv \{p_{odg}\}$ can be summarized as

$$\bar{L}_o = L_o(w;\tau), \text{ for all } o, \tag{1}$$

$$p_{odg} = w_o / \varphi_{odg}, \text{ for all } o, d, \text{ and } g, \tag{2}$$

where φ_{odg} is a productivity shifter and $L_o(w;\tau)$ is the global demand for factor services from country o, which depends on both factor prices and tariffs. Global factor demand can be constructed from primitive assumptions on preferences and technology, as discussed in Adao et al. (2017). For now, it suffices to note that changes in tariffs shift the demand for factor services around the world, as described by equation (1), and in turn these changes affect the prices at which different countries sell their goods, as described by equation (2).

2.2 The Incidence of Tariffs

The causal question of interest can be stated as follows. Suppose global tariffs change from τ to τ'. Holding everything else constant, how do prices adjust, and what are the resulting effects on terms of trade and transfers across countries?

Prices. Consider first a small change from τ to $\tau + d\tau$. The impact of this change on wages can be obtained by differentiating the labor market clearing condition (1). By the Implicit Function Theorem, the changes in wages, $d \ln w$, caused by the changes in tariffs, $d \ln \tau$, are equal to

$$d \ln w = \left(D_{\ln w} \ln L\right)^{-1} \cdot \left(D_{\ln \tau} \ln L\right) \cdot d \ln \tau, \tag{3}$$

where $D_x Y$ denotes the Jacobian matrix $\{\partial Y_j / \partial x_i\}$, i.e. the partial effect of a change in x_i on an outcome Y_j, evaluated at the initial equilibrium. Equation (3) characterizes the reduced-form response of factor prices to tariff changes. It provides the analytical coun-

terpart to Figure 1a, with the price adjustment now expressed in factor space. Given changes in wages, the changes in good prices can then be obtained by using the zero-profit condition (2),

$$d \ln p_{odg} = d \ln w_o. \tag{4}$$

Together, (3) and (4) fully characterize the incidence of a small tariff change $d\tau$ on prices around the world. The incidence of large changes in tariffs can be characterized in the exact same way by integrating over the full path from τ to τ'.

Terms of Trade and Transfers. To go from these price changes into transfers across countries, define the change in country o's terms of trade as

$$d \ln TOT_o = d \ln P_o^x - d \ln P_o^m, \tag{5}$$

with the export and import price indices P_o^x and P_o^m such that

$$d \ln P_o^x = \sum_{d \neq o,g} \omega_{odg}^x d \ln p_{odg}, \tag{6}$$

$$d \ln P_o^m = \sum_{d \neq o,g} \omega_{dog}^m d \ln p_{dog}, \tag{7}$$

where $\omega_{odg}^x \equiv X_{odg} / (\sum_{j \neq o,k} X_{ojk})$ denotes the share of sales X_{odg} of product g to destination d in country o's total exports and $\omega_{dog}^m \equiv X_{dog} / (\sum_{j \neq o,k} X_{jok})$ denotes the share of sales X_{dog} of product g from country d in its total imports.

The income transfer to country o induced by tariff changes, dT_o, corresponds to the change in the value of what it initially sells to the rest of the world (its exports) minus the change in the value of what it initially buys (its imports). Expressed as a fraction of country o's GDP, it is equal to

$$dT_o = \frac{Trade_o}{GDP_o} \times d \ln TOT_o, \tag{8}$$

where $Trade_o$ is the common value of both country o's exports and imports under balanced trade, i.e. $\sum_{d \neq o,g} X_{odg} = \sum_{d \neq o,g} X_{dog}$. This is the analytical counterpart to Figure 1b. Again the impact of large changes can be computed by integrating from τ to τ'.

2.3 Towards an Empirical Specification

The ideal empirical approach to our causal question would directly estimate the reduced-form relationship between $d \ln TOT_o$ and $d \ln \tau$:

$$d \ln TOT_o = \sum_{i,d,g} \beta_{idg,o} d \ln \tau_{ijd},$$

where each coefficient $\beta_{idg,o}$ denotes the partial effect of a single tariff change. In our setting, this would require estimating close to $183^3 \times 5000 \approx 3 \times 10^{10}$ separate coefficients, which is of course not feasible.

As a practical alternative, we leverage the structure from Section 2.2, without fully specifying or estimating a complete model of the world economy. From equations (4) and (6), we know that the export price index reduces to

$$d \ln P_o^x = d \ln w_o, \tag{9}$$

with the change in factor price given by (3). For our baseline analysis, we approximate $(D_{\ln w} \ln L)^{-1} \cdot (D_{\ln \tau} \ln L)$ by

$$\frac{\partial \ln w_o}{\partial \ln \tau_{odg}} = \beta \omega_{odg}^x \text{ for all } d \neq o \text{ and } g, \tag{10}$$

where β is now a single parameter to be estimated. The rationale for this approximation is that the shift in the demand for factor services from country o induced by a single tariff change, $d \ln \tau_{odg}$, on product g in destination d should be proprtional to the size of country o's exports for this product and destination, hence the adjustment by ω_{odg}^x. The extent to which demand shifts translate into prices then depend on unknown elasticities, which an estimate of β is meant to reveal.

Combining (9) and (10), we can express the change in country o's export prices as

$$d \ln P_o^x = \beta \times \sum_{d \neq o,g} \omega_{odg}^x d \ln \tau_{odg}. \tag{11}$$

This is the basis for our baseline specification. In Section 3, we will use (11), along with observed changes in prices and tariffs, to estimate β over different time horizons.

Finally, to compare the prices of the goods that country o sells to those that it buys, note that equations (4)-(7) imply

$$d \ln P_o^m = \sum_{d \neq o} \omega_{do}^m \times d \ln P_d^x,$$

where $\omega^m_{do} \equiv \sum_g \omega^m_{dog}$ denotes the share of imports from country d in country o. Thus, given knowledge of β, one can further compute the change in country o's terms of trade and associated transfers as,

$$d \ln TOT_o = \beta \times \left[\sum_{d \neq o, g} \omega^x_{odg} d \ln \tau_{odg} - \sum_{d \neq o} \omega^m_{do} \times \left(\sum_{d \neq j, g} \omega^x_{djg} d \ln \tau_{djg} \right) \right], \tag{12}$$

$$dT_o = \beta \times \frac{Trade_o}{GDP_o} \times \left[\sum_{d \neq o, g} \omega^x_{odg} d \ln \tau_{odg} - \sum_{d \neq o} \omega^m_{do} \times \left(\sum_{d \neq j, g} \omega^x_{djg} d \ln \tau_{djg} \right) \right]. \tag{13}$$

Equations (12) and (13) open up the possibility of a direct empirical evaluation of whether foreigners pay for tariffs. We turn to this next.

3 Empirical Strategy

Our goal is to estimate the effect on export prices of exposure to foreign tariffs. This section describes the econometric model, assumptions, and data requirements that we deploy in doing so.

3.1 Method

Following the discussion in Section 2, we seek to estimate the effect of any exporting country o's exposure to foreign tariffs $\text{Tariff}_{o,t+1}$—defined below—on outcomes such as its export price index $P_{o,t}$. We do this via a local projection regression based on the following OLS specification:

$$\Delta_h P_{o,t+h} = \beta^h \Delta_1 \text{Tariff}_{o,t+1} + \gamma^h \cdot \text{Controls}_{o,t} + \eta^h_{o,t}. \tag{14}$$

Here, for any variable y_t, the operator $\Delta_h y_{t+h} \equiv y_{t+h} - y_t$ denotes the h-period change. That is, the coefficient β^h in equation (14) represents the effect of an annual change in tariff exposure on export prices over the same year (if $h = 1$), over many subsequent years (if $h > 1$), or even over previous years (if $h < 1$). We present estimates of these coefficients below for many such values of h. Our regression also includes a set of controls, with coefficients estimated separately for each horizon h. The baseline set of variables in the vector $\text{Controls}_{o,t}$ includes country and year fixed effects and several lags of the tariff exposure change variable (e.g. $\Delta_1 \text{Tariff}_{o,t}$ and $\Delta_1 \text{Tariff}_{o,t-1}$).

Exposure to foreign tariffs. Following the logic of equation (11), the one-year change in tariff exposure is defined as a weighted average of (log) tariff changes

$$\Delta_1 \text{Tariff}_{o,t+1} = \sum_{d,g \in \mathcal{S}_{o,t}} \omega^{\tau}_{odgt} \Delta_1 \ln \tau_{odg,t+1}, \tag{15}$$

where we follow the convention that if $t_{odg,t}$ denotes the ad-valorem tariff then $\tau_{odg,t} \equiv 1 + t_{odg,t}$. Letting $X_{odg,t}$ denote the value of such exports from country o to destination market d in product g at time t, we define the weights in this exposure measure as $\omega^{\tau}_{odgt} \equiv X_{odg,t} / \sum_{d,g \in \mathcal{S}_{o,t}} X_{odg,t}$. As discussed further below, the summation in (15) is restricted to a set, denoted $\mathcal{S}_{o,t}$, of foreign tariff changes, in each destination country d and product g, that we believe to be plausibly exogenous from the perspective of exporter o at time t.

Export price index. In order to construct the export price index $P_{o,t}$ we begin with data on all export prices $p_{odg,t}$—that is, the price charged by the exporter exclusive of any import tariff—from origin o, for sale in destination d, of product g, in year t. We then purge these prices of all components that relate to the destination market by projecting $\ln p_{odg,t}$ on a destination-product-year fixed effect and extracting the residual from this regression, denoted $\left(\ln p_{odg,t}\right)^{\text{resid.}}$. This step improves the precision of our results but, as we discuss below, it has very little effect on our point estimates. Finally, we construct the export price index as

$$\Delta_h P_{o,t+h} = \sum_{d,g} \omega^{x}_{odg,t} \Delta_h \left(\ln p_{odg,t+h}\right)^{\text{resid.}}, \tag{16}$$

where $\omega^{x}_{odgt} \equiv X_{odg,t} / \sum_{d,g} X_{odg,t}$. That is, the ($h$-horizon) change in the export price index for country o starting at year t is a sales-weighted average of residualized (h-horizon) changes in the log prices of the country's exports, with weights held fixed in the base year t for all horizons. Notably, all destinations d and products g enter the summation in (16), whereas only the set $\mathcal{S}_{o,t}$ enter the summation in (15).

Identification strategy. The specification in equation (14) relies on a regressor, tariff exposure $\Delta_1 \text{Tariff}_{o,t+1}$, that takes a "shift-share" form—in which the shifters are worldwide tariff changes $\Delta_1 \ln \tau_{odg,t+1}$ and the shares are initial-period export weights $\omega^{\tau}_{odg,t}$. Following Adao et al. (2019) and Borusyak et al. (2022), the OLS estimate of β^h will be unbiased in such a specification when two conditions hold. First, the sum of the shares within any cross-sectional observation o needs to be included in the the vector $\text{Controls}_{o,t}$, separately for each time period t (and horizon h). This is achieved in our case, where

$\sum_{d,g\in\mathcal{S}_{o,t}} \omega^{\tau}_{odg,t} = 1$ for all o and t, through the inclusion of time fixed effects (in each separate horizon-level regression).

The second condition for unbiasedness is that, conditional on the controls included, the shifters need to be as good as randomly assigned relative to the error term $\eta^{h}_{o,t}$. A potential threat to this condition is that country o's trading partners d may raise their tariffs on o precisely when they are experiencing more competition from o's exports, and hence lower export prices embodied in $\eta^{h}_{o,t}$. Following a similar strategy to that in Boehm et al. (2023), we mitigate this concern by focusing on a subset of the data for which we expect that country o's export conditions have negligible effects on the trade policy of partners d.[3]

This restriction works on two levels. First, we isolate cases in which the destination d is applying its MFN tariff $\tau^{MFN}_{dg,t}$ to imports from origin o in product g. Specifically, for any exporter o and year t we define its set of MFN partners $\mathcal{S}^{MFN}_{o,t}$ as the destinations such that the exporter o is subject to MFN status in years t and $t+1$:

$$\mathcal{S}^{\text{MFN}}_{o,t} = \left\{ d, g \text{ such that } \tau_{odg,t} = \tau^{MFN}_{dg,t} \text{ and } \tau_{odg,t+1} = \tau^{MFN}_{dg,t+1} \right\}.$$

This focus on MFN tariffs removes all tariff variation coming from preferential agreements that country d may have with selected partners.

As a second restriction, we further isolate cases in which the exporter o makes a relatively small contribution to country d's imports (among all of the exporters that get MFN treatment from d). Given that MFN trade policies, by definition, apply equally to an entire set of exporters, it is plausible that any endogenous component of d's trade policy is relatively focused on larger MFN partners, who are not included in our analysis. To implement this we require: (i) that exporter o is not in the top 5 MFN partners by trade value for the importer within the 4-digit HS category containing product g, denoted $H^4(g)$; and (ii) that exporter o is not in the top 10 MFN partners by total trade volume for the importer d. Formally, we let $X_{odk,t}$ denote trade values in any product group k and $X_{od,t} \equiv \sum_k X_{odk,t}$. Then, for any exporter o and year t we define its set of small MFN partners as

[3] Boehm et al. (2023) estimated dynamic bilateral trade elasticity specifications (regressions of $\Delta_h \ln X_{odg,t+h}$ on $\Delta_1 \ln \tau_{odg,t+1}$). Their identification strategy was two-fold. First, they dropped from the estimation sample all observations for which (a) the MFN tariff applies, (b) the exporter is "large" in the given destination-product, and (c) the exporter is "large" in the destination as a whole. Second, on the remaining sample they used as an instrument for $\Delta_1 \ln \tau_{odg,t+1}$ this same variable interacted with a dummy for the case in which the tariff $\tau_{odg,t}$ was applying MFN in both t and $t+1$. The reduced-form of this IV procedure amounts to comparing exports, from any given origin, to destinations that charge MFN but for which the exporter is small to destinations that charge PTA tariffs.

$$\mathcal{S}_{o,t}^{\text{Small MFN}} = \left\{ d, g \in \mathcal{S}_{o,t}^{\text{MFN}} \text{ such that } \left\{ \begin{array}{l} X_{odH^4(g),t-1} < X_{dH^4(g),t-1}^{\text{rank=5}},\ X_{odH^4(g),t} < X_{dH^4(g),t}^{\text{rank=5}} \\ X_{od,t-1} < X_{d,t-1}^{\text{rank=10}}, \text{and } X_{od,t} < X_{d,t}^{\text{rank=10}} \end{array} \right. \right\},$$

where $X_{dH^4(g),t}^{\text{rank=5}}$ is the fifth-largest value of d's bilateral imports in $H^4(g)$ in year t and $X_{d,t}^{\text{rank=10}}$ is the equivalent for the tenth-largest value of total bilateral imports. That is, we impose conditions (i) and (ii) in both year t and $t-1$.[4]

Ultimately, with one minor refinement discussed next, the sample of foreign tariff changes $\mathcal{S}_{o,t}$ that we include in the construction of tariff exposure is the "small MFN" subsample, $\mathcal{S}_{o,t}^{\text{Small MFN}}$. This means that estimates of β^h from equation (14) correspond to the causal effect of changes in importers' MFN tariffs on the exporting countries that are "small" from the perspective of the importer in question. Since the tariff variation in the set $\mathcal{S}_{o,t}^{\text{Small MFN}}$ consists of only a subset of the overall tariff variation, β^h can be thought of as the reduced-form effect of an instrumental variable $\Delta_1\text{Tariff}_{o,t+1}$ on the outcome of interest $P_{o,t}$. We will focus on this reduced-form but also discuss the extent to which $\Delta_1\text{Tariff}_{o,t+1}$ appears to affect a version of tariff exposure constructed as per equation (15) but using all tariff variation. This is analogous to a "first stage" relationship between an instrument (the set of exogenous tariff changes) and an endogenous regressor of interest (the set of all tariff changes). We do not emphasize the corresponding IV estimates (i.e. the ratio of the reduced-form β^h to the first stage coefficient) but we discuss below the magnitude of the first stage relationship in order to facilitate interpretation of estimates of β^h.

Further estimation details. Our baseline estimation procedure involves four further details. First, to mitigate concerns about export price outliers, we omit any exporter-importer-product (odg) triplet whose time-series of observations t involves at least one case where the price in adjacent years differs by more than a factor of seven.[5] We apply this restriction to the tariff exposure measure (15) such that the sample $\mathcal{S}_{o,t}$ in (15) consists of the members of $\mathcal{S}_{o,t}^{\text{Small MFN}}$ remaining after this removal of outliers. We also apply this restriction to the sum in the definition of the export price index $P_{o,t}$ (i.e. equation 16). Second, we estimate our regressions on the subsample of country-year pairs ot for which we observe data on export price changes for all horizons $h \in \{1, ..., 5\}$. This ensures that

[4] Classifying countries as small based on $t-1$ expenditures ensures the sample is not selected based on endogenous outcomes.

[5] We choose this value because it corresponds to the largest two percent of maximal within-series price changes across all odg series. We obtain similar results with alternative cutoffs.

each horizon h of equation (14) is estimated on the same sample of country-year pairs, and implies that our main estimates use tariff changes that occur in the 20 years between 1997 and 2016. Third, we estimate specification (14) using regressions that are weighted by the total size (measured with 1995 population) of the exporter o. Finally, we adjust standard errors for clustering by exporter.

3.2 Data Sources

Estimating the coefficient β^h in equation (14) requires three types of data. First, we require data on export prices $p_{odg,t}$ for exporter o at each year t, broken down by their destination d and product g. We use the CEPII BACI database (based on underlying UN COMTRADE records) to obtain this data for essentially all pairs of countries around the world, spanning the period 1995 through 2021, with product categories at the 6-digit HS (1992) level. This database records trade flow values $X_{odg,t}$ in FOB terms, measured in current US dollars calculated at market exchange rates. It also measures trade flow quantities $Q_{odg,t}$ in metric tonnes. We then define the export price as the corresponding unit value, $p_{odg,t} \equiv X_{odg,t}/Q_{odg,t}$.

Second, we require data on the corresponding export share value weights $\omega^x_{odg,t}$ that appear in the export price index $P_{o,t}$ and the analogous weights $\omega^\tau_{odg,t}$ in the tariff exposure measure $\Delta_1 \text{Tariff}_{o,t+1}$. We construct these using the data on $X_{odg,t}$ obtained from the BACI database as well.

Finally, we require data on the ad-valorem tariffs $t_{odg,t} \equiv \tau_{odg,t} - 1$ that each importer d charges—whether via the MFN principle or a preferential agreement—on product g from origin o at time t. We use the Teti (2024) database for this purpose, and omit the small share of cases where reported tariffs are negative or greater than 150%.

In addition, some of the auxiliary regressions reported below explore effects of tariffs on prices, measured in the exporter's currency, and on exchange rates. For this analysis we draw on exchange rate data from the BLS, which is provided at a monthly frequency. We measure annual US dollar exchange rates $\mathcal{E}^{\text{US}}_{o,t}$ as the simple average of the monthly local currency price of US dollars. We convert US dollar export prices to producer currency export prices by applying the annual exchange rate between the exporter's currency and the US dollar, $p^{\text{PCP}}_{odg,t} = \mathcal{E}^{\text{US}}_{o,t} p^{\text{USD}}_{odg,t}$. We then compute a trade-weighted measure of exchange rate depreciation $\Delta_1 \tilde{e}_{o,t+1}$ calculated as the difference between our export price index calculated on the basis of export prices measured in producer currency, $\Delta_1 P^{\text{PCP}}_{o,t+1}$, and that calculated using prices measured in US dollars, $\Delta_1 P^{\text{USD}}_{o,t+1}$. That is,

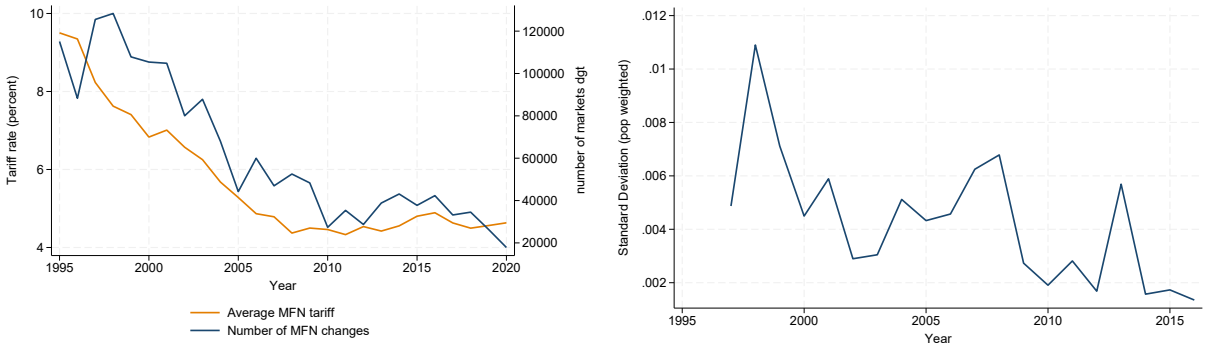
$$\Delta_1 \tilde{e}_{o,t+1} \equiv \Delta_1 P^{\text{PCP}}_{o,t+1} - \Delta_1 P^{\text{USD}}_{o,t+1}.$$

In appendix A we show that this measure corresponds to a trade-weighted exchange rate that accounts for competition among exporters in different import markets.

3.3 Descriptive Statistics

Figure 2: MFN tariff variation over time

(a) MFN tariffs (b) Variation in small MFN exposure



Panel (a) displays the annual average effective MFN tariff alongside the number of MFN tariff changes observed in a given year. The effective tariff rate is defined as the expenditure weighted average tariff among MFN product lines. Panel (b) plots the standard deviation of $\Delta_1 \text{Tariff}_{o,t+1}$ by year t.

MFN tariff variation. Given our focus on the effects of MFN tariff changes, and on the effects of such tariff change on exporters that are "small" with respect to the tariff-setting importers, it is useful to clarify where this variation occurs. To explore this, Panel (a) of Figure 2 plots the annual average of $\tau_{dg,t}^{MFN} - 1$ over time, along with the total number of changes in $\tau_{dg,t}^{MFN}$ (across all importers d and products g) within each year t. A clear message is that the bulk of MFN tariff changes in our 1995-2021 sample took place during the first fifteen years, 1995-2010. Consistent with this, Panel (b) of Figure 2 plots the standard deviation (across all importers d and products g) of our tariff exposure measure $\Delta_1 \text{Tariff}_{o,t+1}$ across exporters o, separately for each year t in our estimation sample. This tells a more nuanced story. The MFN tariff variation during 1995-2010 is certainly apparent, even when focusing only on the subset of variation that hits the exporters that are "small" in the those destination markets that are changing their MFN tariffs. But there is also substantial variation in $\Delta_1 \text{Tariff}_{o,t+1}$ in later years that feeds into the estimates of β^h that we report below.

The dynamics of tariff exposure. Our interest lies in estimating the coefficient β^h in specification (14) for a range of horizons h. If typical changes in ("small MFN") tariffs

that lead to a relatively large value of $\Delta_1 \text{Tariff}_{o,t+1}$ were always perfectly persistent (i.e. such that $\Delta_1 \text{Tariff}_{o,t+1} = \Delta_p \text{Tariff}_{o,t+p}$ for $p > 1$) then the estimates of β^h would reveal the impact of a one-off and permanent tariff change at various ex-post horizons $h > 0$—with increasing values of h corresponding to notions, for instance, of the short, medium and long run of the one-off shock. On the other hand, if the typical tariff change event in our sample is not very persistent then estimates of β^h would correspond to a mix of the impact of a hypothetical permanent tariff shock and the evolving impacts of any subsequent changes in $\Delta_1 \text{Tariff}_{o,t+j}$ in later dates with $j > 1$.

To shed light on this, Figure 3 reports the results from OLS regressions analogous to (14) but with $\Delta_h \text{Tariff}_{o,t+h}$ as the dependent variable. Here, the coefficient at $h = 1$ is mechanically equal to one, while lag controls implies the coefficient at $h = -1$ is mechanically equal to zero. But the estimated coefficients for $h > 1$ reveal the extent to which $\Delta_1 \text{Tariff}_{o,t+1}$ tends to be persistent. If these estimates were equal to one, like the $h = 1$ coefficient is, then we would conclude that $\Delta_1 \text{Tariff}_{o,t+1}$ is perfectly persistent.

The estimates in Figure 3 show that actual "small MFN" tariff changes display modest amounts of mean-reversion, but are not very far from the perfectly persistent case. For example, even at a five-year horizon ($h = 5$) we see an estimated coefficient of approximately 0.7, implying that even five-year persistence is about 70% complete. In addition, this figure suggests that the first few years after a tariff change are where the subsequent changes, if any, occur. This finding is behind our decision to control for two years of lagged changes in $\Delta_h \text{Tariff}_{o,t+h}$, so as to come closer to mimicking a perfectly persistent case. Broadly speaking, the results of Figure 3 suggest that, to a reasonable degree of approximation, the tariff changes that we exploit in studying the effects of exposure in $\Delta_1 \text{Tariff}_{o,t+1}$ are close to perfectly persistent.[6]

4 Do Exporters Pay for Import Tariffs?

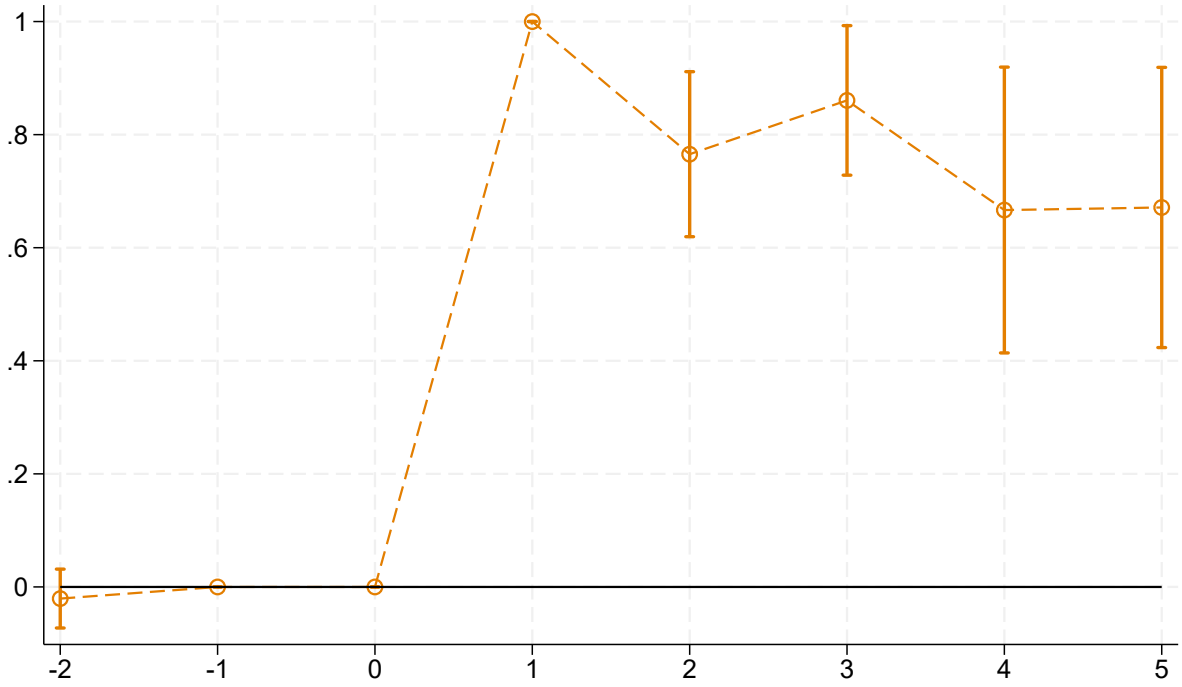
4.1 Impact of Exposure to Foreign Tariffs on Export Price Index

Baseline estimates. We begin by documenting the impact of exposure to foreign tariff increases on export prices across countries. Figure 4 reports estimates of β^h in equation (14) across horizons $h \in \{-2, \ldots, 5\}$, along with the associated 95% confidence intervals.

Estimates indicate that an increase in foreign import tariffs faced by a country leads to a gradual decline in its export prices, with the adjustment unfolding over approximately

[6] In Appendix Figure B.1 we report estimates of the autocorrelation without lagged tariffs as controls, which are again consistent with this conclusion.

Figure 3: Persistence in Exposure Foreign Tariff Changes

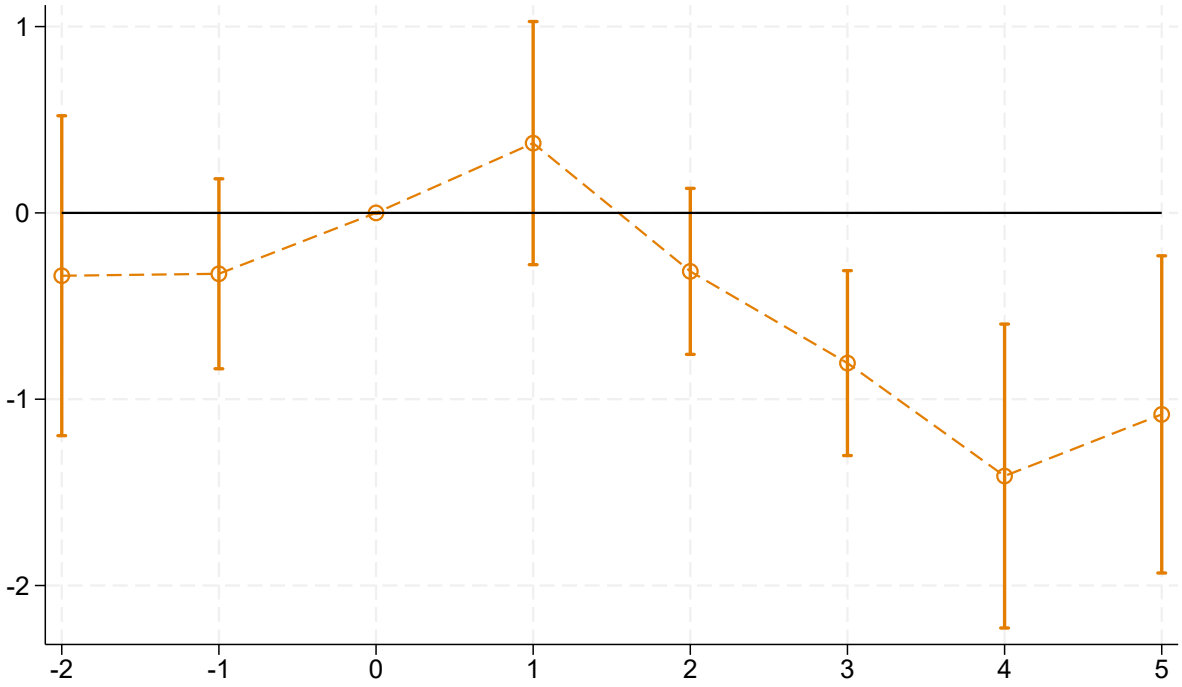


This figures reports estimates of a specification analogous to that in equation (14) but where the dependent variable is $\Delta_h \text{Tariff}_{o,t+h}$. Observations are weighted by 1995 population. Vertical bars correspond to 95% confidence intervals implied by standard errors clustered by country.

four years. Importantly, we find no evidence of pre-trends: export prices do not respond prior to the tariff change. In panel (a) of appendix figure B.3 we report the same estimates, where tariff exposure $\Delta_h \text{Tariff}_{o,t+h}$ is calculated over all trading partners. This regression, using essentially the universe of tariffs changes observed between 1995 and 2017, also shows a similar negative impact unfolding over four years. There is some evidence of pre-trends when using all tariffs, which further motivates our identification strategy.

The estimated impact is economically large. For a country that experiences an increase in the average foreign MFN tariff it faces of 1 percentage point, we estimate a decline in that country's average export prices of 1.4 percentage point after four years. When interpreted as the reduced-form of an IV specification, these magnitudes are even larger. The first-stage coefficient of small MFN tariff changes on the aggregate tariff change is around one half, which implies the magnitudes should be doubled when interpreted as an IV.

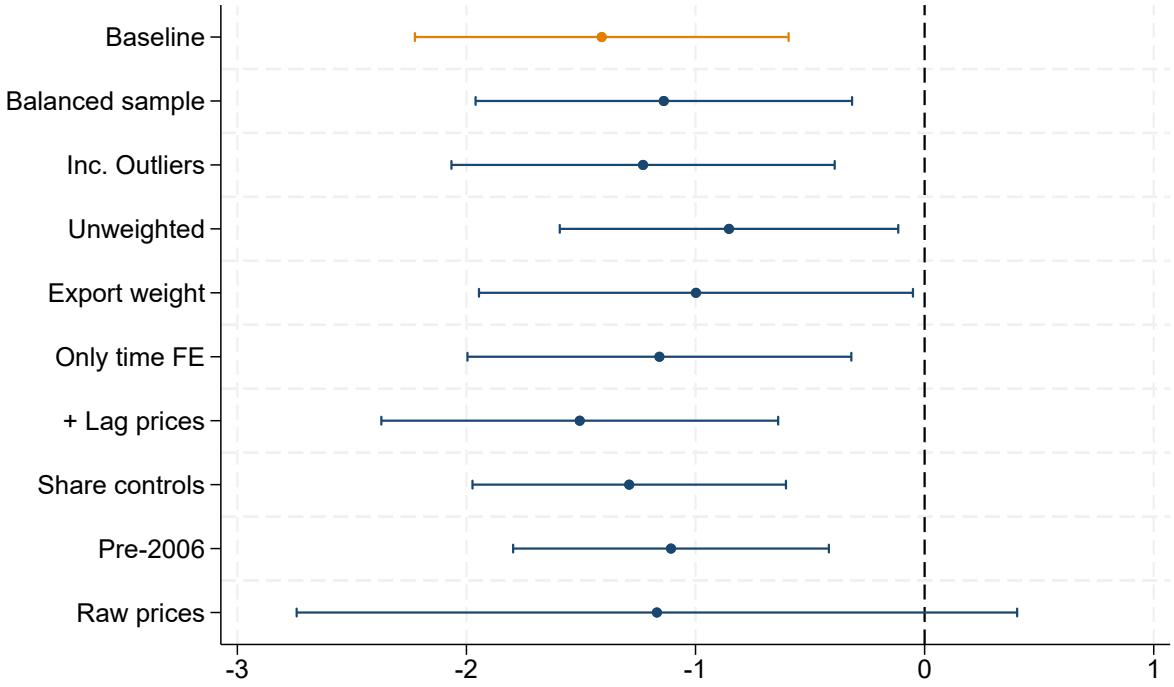
Figure 4: Impact of Exposure to Foreign Tariffs on Export Price Index



Note. This figures reports estimates β^h from equation (14) separately for each horizon $h \in \{-2, ..., 5\}$. We estimate the regression at each horizon h on the subset of country-year pairs for which price changes are available for each $h \in \{1, ..., 5\}$, ensuring a balanced sample of country-year pairs across horizons. Our baseline sample consists of 3,651 country-year pairs, spanning 183 countries between 1997 and 2016. The control vector includes year and country fixed effects as well as two lags of the average foreign tariff. Observations are weighted by each country's population in 1995. Vertical bars correspond to 95% confidence intervals implied by standard errors clustered by country.

Robustness. Figure 5 assesses the robustness of our baseline estimates. For a range of alternative specifications, it reports the estimated effect of changes in average foreign tariffs on the export price index at horizon four, i.e., β^4 from equation (14), together with the corresponding 95% confidence intervals. To facilitate comparison, the first row reproduces the baseline estimate from Figure 4. Each subsequent row presents estimates from specifications that deviate from the baseline along a single dimension, as indicated by the corresponding label.

Figure 5: Impact of Exposure to Foreign Tariffs on Export Price Index, Robustness



Note. This figures reports estimates β^h from equation (14) for horizon $h = 4$ in our baseline sample of country-year pairs. We estimate all regressions on the subset of country-year pairs for which price changes are available for each $h \in \{1, ..., 5\}$. We estimate the regression at each horizon h on the subset of country-year pairs for which price changes are available for each $h \in \{1, ..., 5\}$, ensuring a balanced sample of country-year pairs across horizons. The first row reproduces the baseline estimate from Figure 4. Each subsequent row presents estimates from specifications that deviate from the baseline along a single dimension, as indicated by the corresponding label. "Balanced" reports estimates where $\Delta_4 P_{o,t+4}$ is calculated on the subset of $odgt$ observations for which prices are observed in each year between t and $t + 5$. "Inc. Outliers" reports estimates where all prices are used to compute export price indices, including extreme values of export price changes. "Unweighted" reports estimates where each exporter is weighted equally, excluding countries with less than one million residents or with fewer than 250 observations used to construct tariff exposure. "Export weight" reports estimates where each country is weighted by total exports at year t. "Only time fixed effect" modifies the control set to only include year fixed effects. "Add lagged prices" reports estimates where two lags of one-year export price changes are included as controls."Add share controls" reports estimates in which we control for the share of exports in GDP and the share of BACI exports that are small MFN interacted with year dummies. "Pre-2006" reports the coefficient estimated on the sample of years before 2006. "Raw prices" reports estimates in which export prices are not residualized by market dgt fixed effects when creating an export price index. Observations are weighted by population in 1995, unless stated otherwise. Horizontal bars correspond to 95% confidence intervals implied by standard errors clustered by country.

The first set of alternative specifications modifies the sample of varieties used to construct the export price index, $\Delta_4 P_{o,t+4}$. In the row "Balanced sample," we show that the estimated response is similar when we compute average export price changes using only varieties that exhibit positive trade in every year between t and $t + 5$. Appendix Figure B.2 further shows that restricting the sample in this way does not alter the dynamic

pattern of responses across horizons. This suggests that selection into exporting is not a primary driver of the dynamics documented in Figure B.2. Instead, the decline in export prices reflects adjustments within continuing exporter-destination-product relationships. The row "Inc. Outliers" reports estimates using a version of the export price index that retains extreme changes in unit export values, yielding results similar to the baseline. This indicates that the estimated effects are not affected by a small number of extreme price movements.

The second set of robustness exercises examines the role of weighting across countries. Our baseline specification weights countries by their initial population, assigning greater weight to larger economies. In the row "Unweighted," we show that estimates are slightly smaller when all countries are equally weighted, while excluding countries with fewer than one million residents or fewer than 250 observations used to construct tariff exposure.[7] The row "Export weights" shows that our estimates are also similar when weighting observations by countries' initial aggregate exports, although precision declines relative to the baseline. Together, these results indicate that our findings are not driven by a particular weighting scheme and reflect a systematic pattern across countries, rather than being dominated by a small number of large exporters.

The third set of specifications examines the robustness of our findings to alternative control sets. The row "Only time fixed-effects" shows that our estimate is essentially unchanged when we exclude both country fixed effects and lagged exposure to foreign tariff changes. As shown in Figure 3, this is consistent with the fact that foreign tariff changes resemble one-time, unanticipated, and persistent shocks. The robustness of the estimates to this specification eases concerns about omitted variable bias from slow-moving country characteristics.

To assess the role of serial correlation in export price dynamics, the row "Add lagged prices" augments the specification with lagged changes in average export prices. The estimated coefficient remains virtually unchanged, indicating that our results are not driven by persistence in price adjustments or by mean reversion in export prices.

Next, the row "Add share controls" expands the control set to include interactions between year fixed effects and two country characteristics: the share of exports in GDP and the share of exports for which the country is a small exporter facing MFN tariffs. The similarity of the estimates in this case suggests that our results are not driven by global shocks that differentially affect countries based on their degree of openness or their relative size in destination markets. Instead, the estimated price response seems driven

[7] These estimates remain broadly similar, albeit substantially less precise, when these exclusion criteria are relaxed and smaller countries with limited data are included.

by variation in MFN tariff changes in the destinations of a country's exports.

We next show (in the "Pre-2006" row) that our estimates are similar when we restrict the sample period to between 1995 and 2005, which covers the implementation of the tariff reduction commitments agreed in the Uruguay Round. As shown in Figure 2, this is the period that exhibits most of the variation in exposure to foreign tariff changes. When we instead restrict the sample to the latter period of 2005–2020, we obtain a similar point estimate. However, its standard error is about twice as large as in the full sample, reflecting the lower exposure variation in recent years.

Our final robustness exercise ("Raw prices") considers an alternative measure of export price changes. Our baseline estimates rely on export prices that have been residualized with respect to destination-product-time fixed effects. This removes the direct impact of MFN tariff changes on prices, as well as any price component associated with shocks that are specific to destinations and products. When we instead consider raw price changes, without any residualizing, we obtain a similar point estimate but a wider confidence interval. This indicates that our residualized price measure helps remove noise from export price changes, leading to more precise estimates of the impact of average foreign tariff changes on a country's export price index.

Taken together, these results provide robust evidence that increases in foreign import tariffs lead to economically meaningful declines in export prices that build over time. The effects are not driven by sample composition, weighting choices, or specification details. The consistency of the estimates across a wide range of exercises suggests that the observed price response reflects a systematic adjustment to changes in foreign tariff exposure. In the next section, we turn to the mechanisms underlying these estimated responses in export prices.

4.2 Foreign Tariffs and Export Prices: Micro and Macro Exposure

We have documented economically large responses of export price indices to foreign tariff changes across exporter countries. These findings stand in contrast to estimates in the recent literature that studies the impact of import tariffs on exporter prices at the product level. This literature has focused predominantly on changes in U.S. import tariffs—particularly during the 2018–2019 U.S.-China trade war—and finds relatively small effects on the prices received by exporters (Amiti et al., 2019; Fajgelbaum et al., 2020; Cavallo et al., 2021; Gopinath and Neiman, 2026). This raises a natural question: why do we find large effects when aggregating up to country-level variation in tariff exposure, while studies using product-level variation tend to find small responses?

We address this question using a specification that nests both approaches. Specifically, we estimate the following equation relating changes in micro-level export prices $p_{odg,t}$ to different measures of exposure to foreign tariff changes across origin-good-destination triplets:

$$\Delta_4 \ln p_{odg,t+4} = \beta \Delta_1 \text{Tariff}_{o,t+1} + \beta_{\text{section}} \Delta_1 \text{Tariff}_{os,t+1} + \beta_{\text{variety}} \Delta_1 \text{Tariff}_{dg,t+1} + \gamma \cdot \text{Controls}_{ogd,t} + \eta_{odg,t}, \tag{17}$$

where $\text{Controls}_{ogd,t}$ is a vector of control variables that includes exporter fixed effects and two lags of the exporter's exposure to foreign tariff changes, as well as product-destination-year fixed effects, with products defined at either the four- or six-digit HS level depending on the specification. In line with the logic behind the identification strategy underpinning our baseline estimates, we follow Boehm et al. (2023) by considering only varieties for which the origin country faces MFN tariff changes and accounts for a small share of imports in the destination-good-year. This ensures that our sample only contains varieties whose price changes are unlikely to be important determinants of MFN tariff changes in any given destination.

This specification allows us to decompose price responses across different levels of variation in tariff exposure. As in our baseline macro-level specification, the coefficient β captures the response to the exporter's overall exposure to foreign tariff changes, $\Delta_1 \text{Tariff}_{o,t+1}$ in (15).[8] The coefficient β_{section} captures within-country responses to average tariff changes across broad groups of goods, $\Delta_1 \text{Tariff}_{os,t+1} \equiv \sum_{d,g \in \mathcal{S}_{os,t}} \omega^{\tau}_{odgt} \Delta_1 \ln \tau_{odg,t+1}$ with $\mathcal{S}_{os,t}$ denoting the subset of varieties in $\mathcal{S}_{o,t}$ that belong to HS section s. In addition, β_{variety} captures responses to MFN tariff changes across goods within narrowly defined product categories in the same destination, $\Delta_1 \text{Tariff}_{dg,t+1}$. As a result, β_{variety} is directly comparable to product-level estimates in the literature, which exploit variation in tariffs across goods within narrowly defined categories, conditional on origin-specific shocks.

Table 1 presents the estimates of equation (17) with different sets of regressors. In column (1), we present the analog of our baseline specification, but now estimated directly from micro-level price changes instead of the country's export price index. Specifically, we estimate equation (17) while restricting $\beta_{\text{section}} = \beta_{\text{variety}} = 0$. We obtain an estimate almost identical to the baseline estimate in Figure 4 for the four-year horizon.

In order to exploit variation in tariff changes across goods within the same four-digit product category for a given destination-year pair, we modify the set of fixed effects in our

[8] The estimand is slightly different, as we only consider price changes among small MFN product lines for a given exporter. The fact that estimates are unchanged adds further support to our interpretation of the effects as an aggregate shift in prices.

specification. Column (2) does this by including category-destination-year fixed effects, with a category defined as a four-digit HS code, instead of the good-destination-year fixed effects in our baseline specification. The estimated coefficients show that this has only a marginal impact on our estimates.

The remaining columns report estimates that incorporate the additional measures of exposure to changes in foreign tariffs from equation (17). The estimated price response to the exporter's overall exposure to foreign tariff changes is not affected by the inclusion of these additional measures. Conditional on the origin's overall exposure, column (3) shows that prices do not respond to section-level exposure to foreign tariff changes. Column (4) reports the response of export prices to foreign tariff changes across goods within the same broader product category. The point estimate indicates that an increase of 1 p.p. in tariffs leads to a price decline of 0.26 p.p. (p-value = 0.16). Column (5) indicates that these micro estimates are similar when we include origin-year fixed effects that absorb any exporter-specific shock, including its overall exposure to foreign tariffs.

The results in columns (3)–(5) bridge our baseline estimates in Figure 4 to those in the literature. The typical specification in the literature relies on tariff variation across varieties, conditional on origin-specific shocks. Our estimates indicate that, within our sample of small exporters facing MFN tariff changes in a destination, these variety-level price responses to tariff changes are small. In fact, our point estimate in column (5) is remarkably similar to the long-run estimate of -0.2 in Gopinath and Neiman (2026) for the impact of U.S. tariff changes on import prices during the 2018–2019 U.S.-China trade war (column 1 of Panel A in Table 1). This indicates that product-level export price responses remain small even outside the U.S. context and over a longer time horizon (1995–2021). However, the macro-level responses operating at the country level are larger in magnitude. This suggests that the adjustment occurs at the country level, not within narrowly defined product categories, as we explore next.

4.3 Foreign Tariffs, Producer Currency Prices, and Exchange Rates

Having established that price adjustments operate at the macro level, we now investigate one key channel: adjustments across currencies. Specifically, we estimate equation (14) using two additional dependent variables defined in Section 3.2: the export price index in the producer's currency ($\Delta^h P^{\text{PCP}}_{o,t+h}$), and the export-weighted exchange rate between the U.S. dollar and the producer's currency ($\Delta_h \tilde{e}_{o,t+h}$).

Figure 6 reports the resulting estimates of β^h across horizons $h \in \{-2, \ldots, 5\}$, along with 95% confidence intervals. Panel (a) shows that export prices in the producer's cur-

Table 1: The Impact of Foreign Tariffs on Export Prices: Micro and Macro

	(1)	(2)	(3)	(4)	(5)
β	-1.41	-1.20	-1.21	-1.21	
	(0.67)	(0.63)	(0.61)	(0.61)	
β_{section}			0.00	0.02	-0.12
			(0.25)	(0.25)	(0.25)
β_{variety}				-0.27	-0.26
				(0.19)	(0.19)
Origin year dummies:	No	No	No	No	Yes
Destination-year dummies interacted with:					
6-digit HS code:	Yes	No	No	No	No
4-digit HS code:	No	Yes	Yes	Yes	Yes

Note. This table reports estimates of equation (17) using a sample of 24,671,062 origin-good-destination-year observations in which the origin country faces MFN tariff changes and accounts for a small share of imports in the destination-good-year. The control vector includes exporter fixed effects and two lags of the origin's exposure to foreign tariff changes, as well as the set of fixed effects described in the bottom rows of each column. All regressions weight observations by initial trade flows. Standard errors, reported in parentheses, are clustered by exporter o.

rency display little response to foreign tariff changes. The point estimates are positive, but small and imprecise. Four years after a 1 percentage point increase in foreign tariff exposure, export prices in the producer's currency rise by 0.3 percentage points, with a standard error of .99 (p-value = .764).

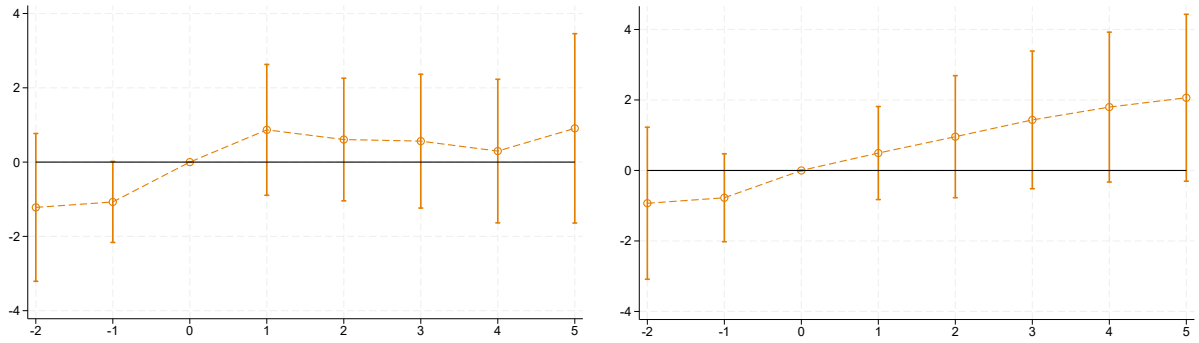
Panel (b) clarifies the adjustment mechanism. Higher foreign tariffs lead to a strong depreciation of the exporter's currency relative to its trading partners, with effects that build over time. A 1 percentage point increase in foreign tariff exposure is associated with a 2 percentage point depreciation after four years. The pair of results in Figure 6 indicate that the adjustment in export prices operates primarily through exchange rate movements rather than changes in producer-currency prices.

5 Concluding Remarks

This paper provides direct empirical evidence on the incidence of import tariffs on exporter countries. Using a difference-in-differences approach across countries and over time, we show that increases in foreign tariffs lead to large and persistent declines in export prices, consistent with a deterioration in the terms of trade of exposed countries.

Figure 6: Impact of Exposure to Foreign Tariffs on Export Prices and Exchange Rates

(a) Producer Currency Price (b) Exchange Rate



Note. This figures reports estimates β^h from equation (14) separately for each horizon $h \in \{-2, ..., 5\}$, the dependent variable is the average export price in the producer currency in Panel (b) and the exchange rate between the U.S. dollar and producer currency in Panel (c) . We estimate the regression at each horizon h on the subset of country-year pairs for which price changes are available for each $h \in \{1, ..., 5\}$, ensuring a balanced sample of country-year pairs across horizons. Our baseline sample consists of 3,651 country-year pairs, spanning 183 countries between 1997 and 2016. The control vector includes year and country fixed effects as well as two lags of the average foreign tariff. Observations are weighted by each country's population in 1995. Vertical bars correspond to 95% confidence intervals implied by standard errors clustered by country.

These aggregate adjustments operate primarily through exchange rates. Export prices in producer currency show little response, while prices in U.S. dollars decline alongside currency depreciations. Taken together, our findings indicate that foreigners do bear a meaningful share of the burden of tariffs. They underscore the importance of studying tariff incidence through macro channels that operate at the country-level.

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A Exchange Rate Definition

Our export price index is defined as

$$\Delta P_o \equiv \sum_d \omega_{od} \left(\Delta p_{od} - \hat{E} \left[\Delta p_{od} | d \right] \right),$$

where $\hat{E}\left[\Delta p_{od}|d\right] = \sum_i \omega_{id} \Delta p_{id}$ is the sample-weighted average within a destination d. We define the exchange rate as

$$\Delta \log \epsilon_o \equiv \Delta P_o^{\text{PCP}} - \Delta P_o^{\text{USD}},$$

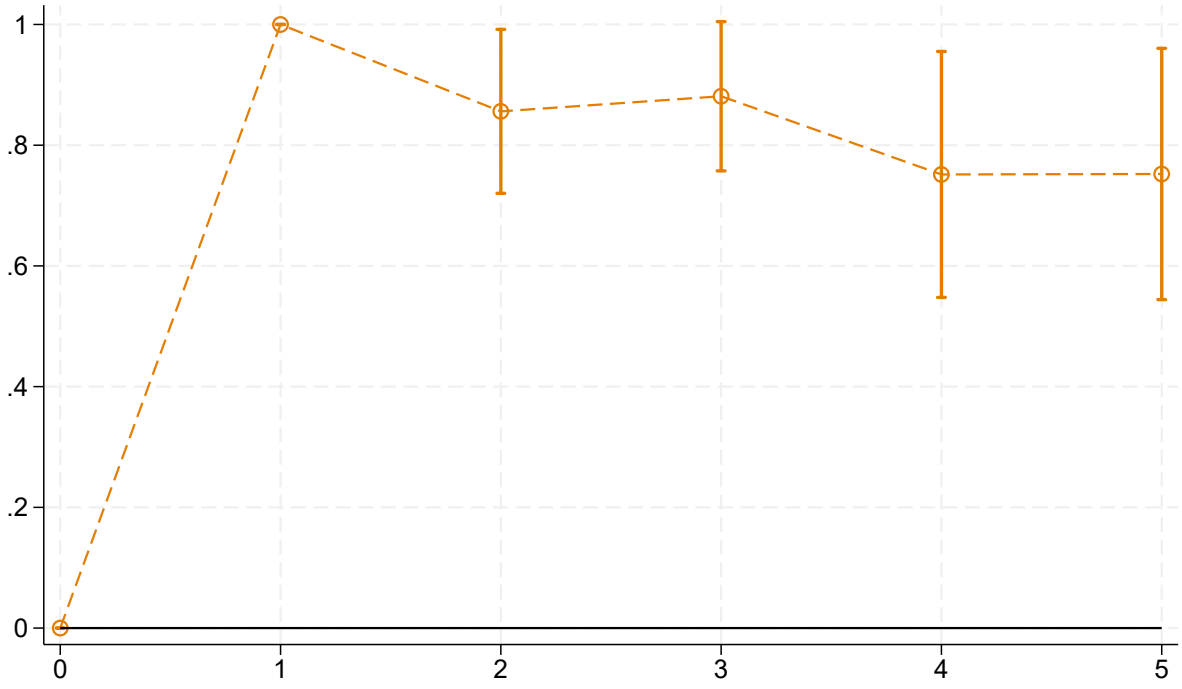
while we know micro prices can be converted according to $\Delta p_{od}^{\text{PCP}} = \Delta \log \epsilon_o^{\text{USD}} + \Delta p_{od}^{\text{USD}}$. Therefore the exchange rate we define is

$$\begin{aligned}
\Delta \log \epsilon_o &= \sum_d \omega_{od} \left(\Delta \log \epsilon_o^{\text{USD}} + \Delta p_{o,p} - \hat{E} \left[\Delta \log \epsilon_i^{\text{USD}} + \Delta p_{id} | d \right] \right) - \sum_d \omega_{od} \left(\Delta p_{od} - \hat{E} \left[\Delta p_{od} | d \right] \right) \\
&= \sum_d \omega_{od} \left(\Delta \log \epsilon_o^{\text{USD}} - \hat{E} \left[\Delta \log \epsilon_i^{\text{USD}} | d \right] \right).
\end{aligned}$$

We can interpret $\hat{E}\left[\Delta \log \epsilon_i^{\text{USD}} | d\right]$ as an import-weighted exchange rate for country d. This means that our exchange rate is an export-weighted exchange rate index, where the exporter exchange rate is measured relative to an import weighted exchange rate for each destination

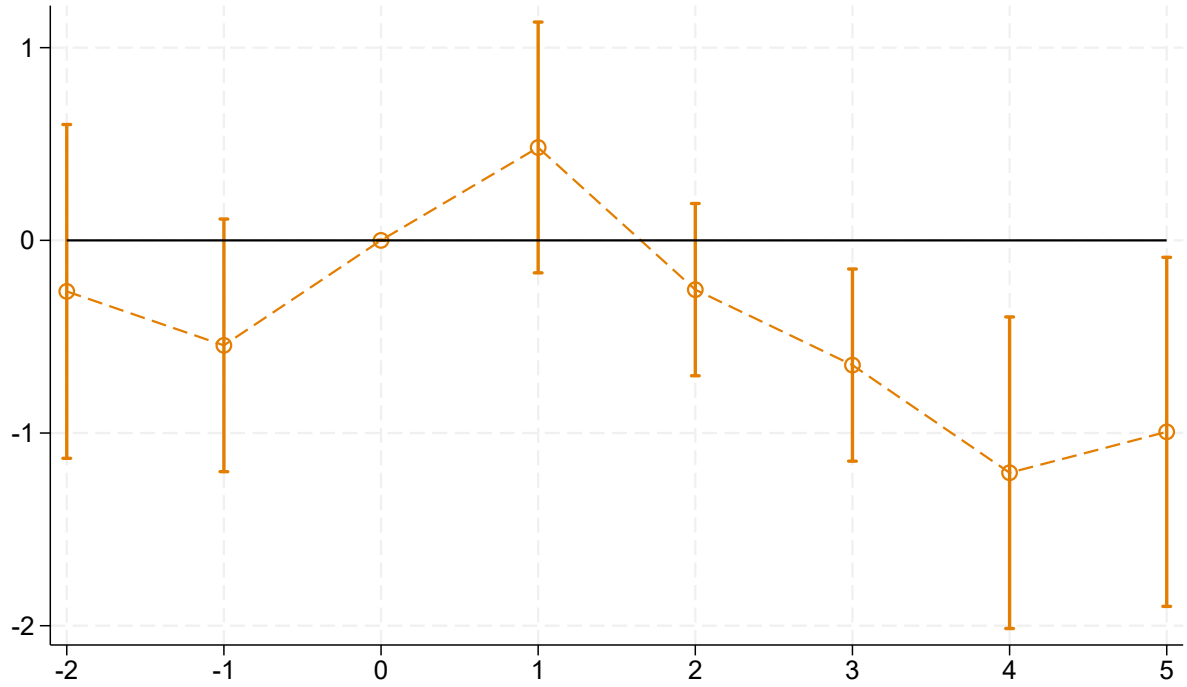
B Additional Empirical Results

Figure B.1: Autocorrelation of Average Foreign Tariff Changes



Note. This figures reports estimates of the simple autocorrelation of tariff changes at different horizons ρ^h, estimated from the regression $\Delta_h \text{Tariff}_{o,t+h} = \delta_o + \delta_t + \rho^h \Delta_1 \text{Tariff}_{o,t+1} + \varepsilon^h_{o,t+h}$ separately for each horizon $h \in \{-2, ..., 5\}$. The dependent variable is the average foreign tariff change between years t and $t+h$, $\Delta_h \text{Tariff}_{o,t+h} \equiv \sum_{d,g \in \mathcal{S}_{o,t}} \omega_{odgt} \Delta_h \ln \tau_{odg,t+h}$. We estimate the regression at each horizon h on the subset of country-year pairs for which tariff changes are available for each $h \in \{1, ..., 5\}$, ensuring a balanced sample of country-year pairs across horizons. Our final estimation consists of 3,645 country-year pairs, spanning 182 countries between 1997 and 2016. Observations are weighted by each country's population in 1995. Vertical bars correspond to 95% confidence intervals implied by standard errors clustered by country.

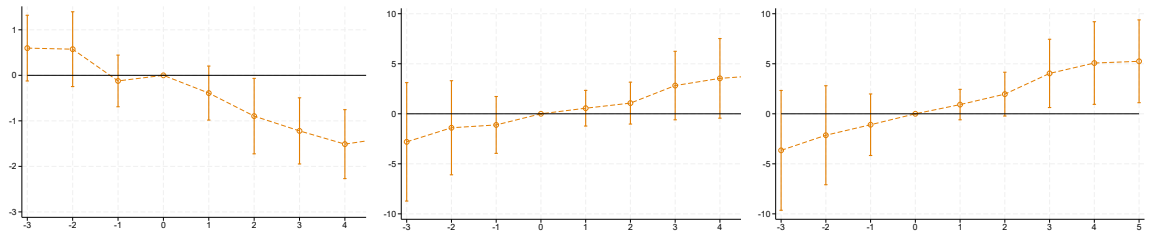
Figure B.2: Impact of Exposure to Foreign Tariffs on Export Price Index, Balanced Sample of Exported Varieties



Note. This figures reports estimates β^h from equation (14) separately for each horizon $h \in \{-2, ..., 5\}$. We estimate the regression at each horizon h on the subset of country-year pairs for which price changes are available for each $h \in \{1, ..., 5\}$, ensuring a balanced sample of country-year pairs across horizons. Our baseline sample consists of 3,642 country-year pairs, spanning 183 countries between 1997 and 2016. The export price index is computed as the export-weighted average price change across varieties with positive trade in every year between t and $t + 5$. The control vector includes year and country fixed effects as well as two lags of the average foreign tariff. Observations are weighted by each country's population in 1995. Vertical bars correspond to 95% confidence intervals implied by standard errors clustered by country.

Figure B.3: Impact of Exposure to Foreign Tariffs on Export Prices and Exchange Rates, Average Foreign Tariff Across All Varieties

(a) U.S. Dollar Price (b) Producer Currency Price (c) Exchange Rate



Note. This figures reports estimates β^h from equation (14) separately for each horizon $h \in \{-2, ..., 5\}$ in which the dependent variable is the average export price in U.S. dollars in Panel (a), average export price in the producer currency in Panel (b), and the exchange rate between the U.S. dollar and producer currency in Panel (c). Average exposure to foreign tariff changes computed across all varieties, $\Delta_1 \text{Tariff}_{o,t+1} = \sum_{d,g \in \mathcal{S}_{o,t}} \omega_{odgt} \Delta_1 \ln \tau_{odg,t+1}$ where $\mathcal{S}_{o,t}$ includes all varieties with data on export price changes. The control vector includes year and country fixed effects as well as two lags of the average foreign tariff. Observations are weighted by each country's population in 1995. Vertical bars correspond to 95% confidence intervals implied by standard errors clustered by country.