The Heterogeneous Impact of the 2018–2019 Trade War on U.S. Prices and Welfare

Sang Hoon Kong*

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Abstract

A puzzling finding from the 2018–2019 trade war is that tariffs by large economies like the U.S. and China seemingly failed to lower pre-tariff prices. In this paper, I show that price reductions did occur but were concentrated in products where importers held significant bilateral market power. Because these products represent a large share of trade value, I estimate a 10% value-weighted average decline in import prices and a 38% pass-through rate for U.S. tariffs. The evidence on export prices is much weaker. As a result, the U.S. experienced a net positive terms-of-trade welfare gain.

JEL Codes: F10, F13, F14

Keywords: trade war, tariff pass-through, market power, trade policy, terms-of-trade, difference-in-differences

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1 Introduction

A central puzzle from the 2018–2019 trade war is that tariffs imposed by large economies such as the U.S. and China seemingly failed to lower exporters' prices. Standard trade theory predicts that tariffs imposed by large importers should reduce foreign export prices and generate terms-of-trade gains. Instead, a robust, yet surprising finding in the literature is that the trade war tariffs had nearly complete pass-through (Amiti et al., 2019; Fajgelbaum et al., 2020; Flaaen et al., 2020; Amiti et al., 2020; Cavallo et al., 2021; Ma et al., 2021; Jiao et al., 2024). Exporters did not lower their prices, and importers ended up bearing the full cost of the tariffs. As a result, trade economists concluded that the cost of U.S. tariffs was almost entirely borne by its own consumers and importers.

In this paper, I revisit the impact of the 2018–2019 trade war on U.S. prices and welfare. I show that import price reductions did occur but were concentrated in products where importers held significant bilateral market power over the exporter. Because these products account for a disproportionately large share of trade value, I estimate a 10.3% value-weighted decline in pre-tariff import prices and an aggregate pass-through rate of 38% for U.S. tariffs by the end of 2019. In contrast, the evidence on pre-tariff export prices and on the pass-through of retaliatory tariffs is much weaker. Thus, the main U.S. price adjustment occurred on the import side.

Combining the estimated effects with a standard framework (Dixit and Norman, 1980), I estimate a U.S. welfare gain of 0.17% of 2017 GDP, or $32.95 billion, by the end of 2019. This gain is driven by a positive net terms-of-trade effect. The positive welfare gain lends empirical support to the optimal tariff literature (Johnson, 1953; Syropoulos, 2002) that argues large importing countries can benefit from tariffs by manipulating their terms of trade.[1] These results overturn the view that the costs of U.S. tariffs were borne almost entirely by U.S. importers through 2019.

I address a couple of key identification issues to arrive at these conclusions. While previous studies have used tariff implementation dates as treatment dates and not-yet-tariffed products as the control group, I show that these choices likely violate standard difference-in-differences (DiD) identification assumptions. Because tariffs were announced in advance and credible threats were made even earlier, I show that using tariff implementation dates as treatment dates is often too late to capture their effects. Furthermore, I argue that not-yet-tariffed products, especially imports from and exports to China, should not be considered untreated since tariffs were announced on them but were later called off. Once these identi-

[1] While Broda et al. (2008) find that inverse export supply elasticities correlate with unilaterally implemented tariffs around the world, they do not directly show that such tariffs lead to terms-of-trade gains.

fication issues are addressed, DiD estimates reveal a sizable decline in value-weighted import prices once the trade war begins but not leading up to it.

Using the imputation-based estimator of Borusyak et al. (2024), I find that the aggregate effects on U.S. prices mask substantial heterogeneity across different waves of tariffs. In particular, there are statistically significant import price declines and evidence of incomplete pass-through only for the third and fourth waves of U.S. tariffs on China. However, the aggregate effect is heavily influenced by the estimates for these two waves as they cover 73% of all tariffed import values.

Finally, I show that bilateral importer market power is central to understanding this heterogeneity. Aggregate importer size is not the right measure of market power in this setting because the trade war was concentrated in bilateral trade flows, especially between the U.S. and China. What matters instead is bilateral importer market power, which measures the extent to which a specific exporter can redirect sales to alternative markets outside the tariff-imposing country. Based on this novel measure, I confirm that both the U.S. and China are large importers but nevertheless have very limited bilateral market power over one another for most products. These results reconcile why unweighted estimates indicate nearly complete pass-through for the U.S. and China. Both countries have little bilateral market power over one another despite their overall large market shares.

Indeed, I find that tariff pass-through rates are systematically lower for higher levels of bilateral importer market power. As a result, incomplete pass-through of both U.S. and foreign tariffs is concentrated in a smaller share of products where they hold greater bilateral market power. For more than half of treated import varieties, including the median variety, I cannot reject complete pass-through. Thus, the aggregate decline in import prices is driven disproportionately by imports with large values, because these are the products in which the U.S. had greater market power over the exporter. The retaliating countries had market power over U.S. exporters in even fewer products, which is consistent with the lack of export price responses.

The seeming inability of the U.S. and China to influence their terms of trade during the 2018–2019 trade war has been puzzling to trade economists (Fajgelbaum and Khandelwal, 2022). Recent work has increasingly turned to firm-level data and richer models to explain this puzzle (Sui et al., 2022; Fan et al., 2025). My findings offer a complementary explanation to the puzzle and should caution future studies of trade wars from overlooking bilateral importer market power and using unweighted estimates.

The paper proceeds as follows. Section 2 describes the trade and tariff data and introduces the bilateral importer market power measure. Section 3 presents the empirical framework, identification issues, and estimation strategy. Section 4 reports the main results on U.S.

prices, tariff pass-through, and welfare. Section 5 studies the systematic relationship between the heterogeneity in tariff pass-through and the bilateral importer market power.

2 Data

This section describes the trade and tariff data, defines the treatment groups, and introduces the market-power measures used later to study heterogeneity in tariff pass-through.

2.1 Trade War Tariffs and Treatment Groups

I use monthly U.S. Census data on imports and exports from 2017 to 2020. Throughout the paper, I use data on "imports for consumption" (as opposed to general imports) and "domestic exports" (as opposed to total exports). I refer to a unique pair of 10-digit product code and trading country as a variety. For each variety, the data include trade value and, when reported, quantity. I measure pre-tariff prices as unit values, computed by dividing the trade value by quantity.[2] For imports, the data also report collected duties, which I use to construct baseline U.S. tariff rates in 2017. The main analysis focuses on 2018–2019, since 2020 is potentially contaminated by pandemic-related shocks, although I report 2020 estimates to assess the stability of my findings.

U.S. trade-war tariffs from 2018–2019 are collected from official U.S. announcements and mapped directly to U.S. import varieties at the 10-digit level. This mapping is precise because the tariff schedules and the import data use the same product classification. In contrast, foreign retaliatory tariffs cannot be matched as precisely. Retaliating countries report tariffs at their own tariff-line level, and product classifications are internationally comparable only up to the first six digits. I therefore aggregate retaliatory tariffs to the 6-digit level before merging them with U.S. export data. To construct total retaliatory tariff rates, I start from the 2017 applied MFN rates from the World Bank's WITS and add the retaliatory tariff increases.[3] Appendix A provides further details on these tariffs.

It is important to note that the retaliatory tariffs are measured less accurately than the import tariffs for two reasons. First, all U.S. export varieties sharing the same first six digits of product code are assigned the same retaliatory tariff rate. Second, the foreign tariff schedules are aggregated using simple averages rather than trade-value weights, because

[2]To guard against extreme measurement errors, I do not use an observation's quantity information (consider it as missing) if the ratio between the unit value and the median unit value for the same variety is greater than 10 or lower than 1/10. I also do not use the quantity information if the measurement unit for the variety's quantity changed during the sample period.

[3]For Canada and Mexico, I use the preferential rates for the U.S. when applicable.

Table 1: Summary Statistics by Treatment Group

Panel A: U.S. Import Tariffs

Treatment Group	Initial Implementation	Import Value (Million USD)	Tariff Rate (%) 2017	2018	2019
Solar/Washer	Feb 07, 2018	9,011	0.5	30.3	27.0
Aluminum/Steel 1	Mar 23, 2018	18,071	0.6	20.6	21.7
Aluminum (KOR)	May 01, 2018	128	0.7	10.7	10.7
Aluminum/Steel 2	Jun 01, 2018	22,912	0.2	19.3	7.6
China 1	Jul 06, 2018	31,864	1.0	26.0	26.0
China 2	Aug 23, 2018	13,259	1.7	26.7	26.7
China 3	Sep 24, 2018	186,138	2.1	12.1	28.9
China 4	Sep 01, 2019	98,884	6.8	6.8	21.8
E.U. Dispute	Oct 18, 2019	10,072	0.8	0.8	18.3
Never Tariffed (China)	∞	168,982	1.3	1.3	1.3
Never Tariffed (ROW)	∞	1,767,832	1.1	1.1	1.1
Total		2,327,153	1.4	3.2	5.1

Panel B: Foreign Tariffs

Treatment Group	Initial Implementation	Export Value (Million USD)	Tariff Rate (%) 2017	2018	2019
Aluminum/Steel: China	Apr 02, 2018	2,443	7.6	52.7	57.5
Aluminum/Steel: Mexico	Jun 05, 2018	4,231	0.0	14.0	0.0
Aluminum/Steel: Turkey	Jun 21, 2018	1,286	12.1	35.8	23.8
Aluminum/Steel: E.U.	Jun 22, 2018	3,788	2.8	22.0	22.0
Aluminum/Steel: Canada	Jul 01, 2018	16,594	0.0	12.4	0.0
China 1	Jul 06, 2018	27,806	12.6	31.5	34.8
Aluminum/Steel: Russia	Aug 05, 2018	259	3.2	16.3	16.3
China 2	Aug 23, 2018	11,661	3.6	26.8	27.6
China 3	Sep 24, 2018	43,669	5.6	12.2	19.6
Aluminum/Steel: India	Jun 16, 2019	1,259	9.0	9.0	22.5
China 4	Sep 01, 2019	5,098	0.7	0.7	5.2
Never Tariffed (China)	∞	29,431	0.0	0.0	0.0
Never Tariffed (ROW)	∞	1,160,868			
Total		1,308,395			

Notes: Trade values are for 2017 based on "Imports: For Consumption (Customs Value)" and "Exports: Domestic" from the U.S. Census Bureau. The symbol ∞ indicates that the implementation of tariffs never occurred during the sample period. The end-of-year tariff rates are calculated by averaging product-country level tariff rates weighted by import or export values in 2017. The foreign tariff rates for 6-digit product codes are calculated by averaging respective trade partner's complete panel of tariff rates at the 8-digit level. The tariff rates on Solar/Washer products are contingent on a quantity threshold.

tariff-line-level import values for the retaliating countries are not readily available. These limitations matter more for estimates that use tariff magnitudes directly, such as retaliatory

tariff pass-through rates, than for non-tariff outcomes such as pre-tariff export prices.

Table 1 summarizes the treatment groups by first implementation date. The table shows that the 2018–2019 trade war was highly concentrated in bilateral U.S.-China trade. In particular, China 3 and China 4 account for 73% of import value exposed to U.S. tariffs, while China 3 alone accounts for 37% of export value exposed to foreign retaliation. These groups therefore play a central role in the aggregate price and welfare effects estimated later in the paper.

I also distinguish never-tariffed Chinese varieties from the never-tariffed varieties from the rest of the world. This is because both U.S. and Chinese tariffs were scheduled to go into effect on December 15, 2019 for the majority of these products. With the agreement over the Phase One trade deal between the two countries, however, these tariffs were called off. Despite never being implemented, they were announced on May 17, 2019 by the U.S. (Federal Register, Vol. 84, No. 96) and August 23, 2019 by China along with detailed product lists. I therefore keep them separate from the broader never-tariffed rest-of-world group and return to the identification implications of this distinction in Section 3.2.

2.2 Importer Market Shares and Market Power

To study heterogeneity in exporters' price responses, I use 2017 CEPII-BACI data at the 6-digit product level (Gaulier and Zignago, 2010). These data are needed because the U.S. Census data do not contain non-U.S. trade flows, and the relevant notions of importer market power depend on a country's importance in global demand.

I begin with a standard measure of importer size. For each 6-digit product, importer market share is defined as the importing country's total imports of that product divided by world imports of the same product:

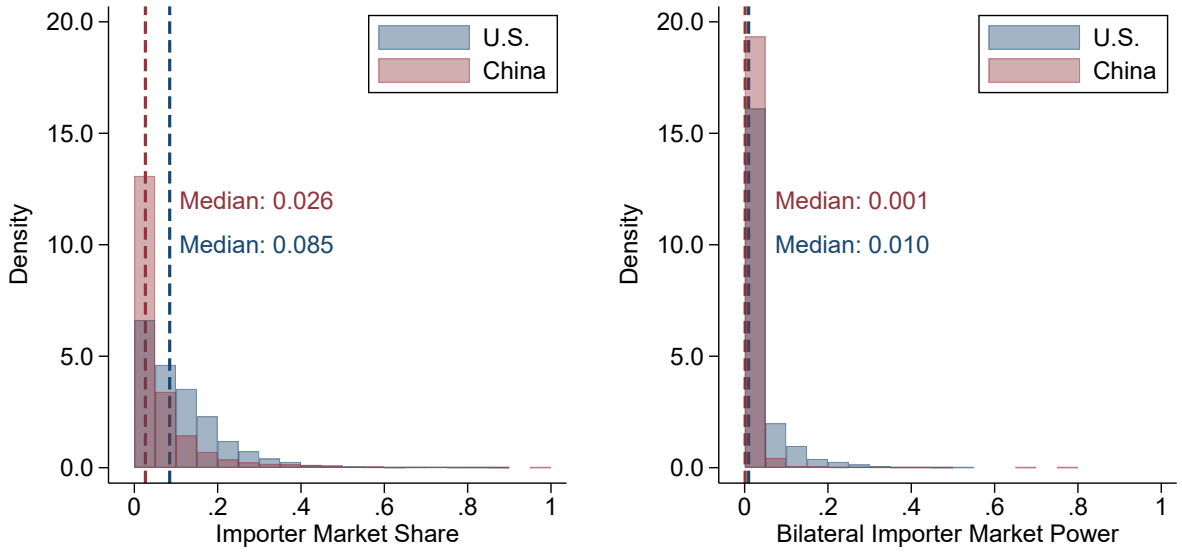
$$\text{Importer Market Share}_{ci} = \frac{\sum_j m_{cij}}{\sum_k \sum_j m_{kij}}, \tag{1}$$

where m_{cij} denotes country c's imports of product i from exporter j.[4]

In the left panel of Figure 1, I plot the distribution of U.S. and China's importer market shares across 6-digit products. The data show that both countries are indeed large importers. The median importer market share is 8.5% for the U.S. and 2.6% for China, and the U.S. exceeds a 10% share for about 44% of products, compared with 18% for China. These facts help explain why trade economists have found previous results on the 2018–2019 trade war puzzling. Despite their size, both U.S. and Chinese tariffs seemingly failed to induce

[4]Due to lack of data, I do not account for domestic consumption.

Figure 1: U.S. and China's Importer Market Shares and Bilateral Importer Market Power



Notes: The left figure plots the distribution of importer market shares for the U.S. and China across 6-digit product codes. Importer market share divides an importer's total value of imports by the world's total value of imports of the same good (equation (1)). The right figure plots the distribution of bilateral importer market power in equation (2) between the two countries across 6-digit product codes. Bilateral importer market power divides an importer's value of imports from an exporter by the world's residual value of import of the same good (equation (2)). These measures are calculated using 2017 international trade data from CEPII-BACI.

exporters to lower their prices.

The importer market share, however, may not capture an importer's market power over a specific exporter. For example, consider a product for which the U.S. has a 90% market share. An exporter that is responsible for 20 percentage points of this market share will find it difficult to find alternative markets to reallocate its sales if tariffs make exporting to the U.S. more costly. However, an exporter that is responsible for only 1 percentage point of this market share will find it easier to reallocate its exports since the remaining market outside the U.S. is large enough to cover their exports. As the 2018–2019 trade war heavily featured tariffs on bilateral trade flows between the U.S. and China, it is important to capture this bilateral aspect of importer market power.

I therefore define bilateral importer market power as

$$\text{Bilateral Importer Market Power}_{cij} = \frac{m_{cij}}{\sum_k \sum_o m_{kio} - \sum_{o \neq j} m_{cio}}. \tag{2}$$

This variable divides importer c's value of imports from country j by the *residual* world import value of other countries and the value of the bilateral trade being considered. In other words, the denominator captures the residual demand that country j faces should it

try to reallocate its exports to other countries, plus the current bilateral trade value. For example, if the importing country has a monopoly over product i, then there is no residual demand and the denominator is exactly equal to the numerator (m_{cij}), implying a bilateral importer market power of 1. In contrast, the importer has lower market power when the bilateral trade flow is small relative to the residual demand. Thus, all else equal, an importer has greater bilateral market power over trade partners that it imports heavily from.

The distributions of U.S. and China's bilateral importer market power over the other are plotted in the right panel of Figure 1. The median bilateral importer market power is much smaller than the aggregate importer shares suggest: about 1% for the U.S. over China and close to zero for China over the U.S. At the same time, the upper tail displays more meaningful market power. The U.S. exceeds 5% bilateral market power for about 19% of products, compared with only 3% for China. These patterns suggest that tariff pass-through should be most incomplete in products where bilateral importer market power is high, a prediction I test in Section 5.

3 Empirical Framework and Estimation

3.1 Target Parameters

To guide the empirical work, I lay down a simple theoretical framework for assessing the welfare impact of trade war tariffs.

Consider an initial equilibrium indexed by $t = 0$ prior to the tariff changes. I assess the welfare impact of tariff changes at time t relative to the initial equilibrium by its equivalent variation, a hypothetical change in domestic income that would have had an equivalent welfare impact as the tariff changes. Dixit and Norman (1980) (also see Fajgelbaum and Khandelwal (2022)) show that the first-order approximation of the equivalent variation at time t can be calculated as

$$EV_t = \sum_{ij \in \Omega^M} r_{ij0} \Delta \ln q^m_{ijt} - \sum_{ij \in \Omega^M} m_{ij0} \Delta \ln p^{m*}_{ijt} + \sum_{ij \in \Omega^X} x_{ij0} \Delta \ln p^{x*}_{ijt}, \tag{3}$$

where m_{ijt} is the import value of product i from country j at time t; x_{ijt} is the value of export; p^{m*}_{ijt} and p^{x*}_{ijt} are the pre-tariff prices of imports and exports, respectively; q^m_{ijt} is the quantity of import; τ^m_{ijt} is the ad-valorem import tariff such that $p^m_{ijt} = (1+\tau^m_{ijt})p^{m*}_{ijt}$ is the final price that importers face; and $r_{ijt} = \tau^m_{ijt} m_{ijt}$ is the tariff revenue. The difference operator Δ is defined with respect to the initial equilibrium in $t = 0$ such that $\Delta \ln p^{m*}_{ijt} = \ln p^{m*}_{ijt} - \ln p^{m*}_{ij0}$, for example. Thus, $\Delta \ln p^{m*}_{ijt}$ measures the effect of tariff changes on the pre-tariff import

price of variety ij at time t.

The first two terms in equation (3) capture the net effect of import tariff changes on domestic consumers, producers, and government revenue. The equation shows that import tariffs affect tariff revenues through changes in import quantities (the first term), the effects of which can be offset if the tariffs induce foreign exporters to lower their pre-tariff prices (the second term). Finally, the third term captures the effect of retaliatory tariffs on domestic exporters, which will be more negative if foreign tariffs induce domestic exporters to lower pre-tariff prices. The net effect on pre-tariff import and export prices (the second and third terms) is often referred to as the terms-of-trade effect.

The equivalent variation in equation (3) can be re-expressed as a function of the average treatment effect on treated units (ATT) specific to each treatment group, which will be more convenient for empirical work. Let Ω^M denote the set of varieties imported by the U.S. Each variety ij is associated with an *initial* date of treatment T_{ij}. Following Baker et al. (2025), I use $T_{ij} = \infty$ to denote that a variety was never treated during the sample period. Eventually treated imported varieties are partitioned into different treatment groups $g \in G^M$ based on their initial date of treatment, where G^M denotes the set of all treatment groups. To be clear, if a variety is subject to multiple rounds of treatment, the treatment-group assignment will only depend on the initial date of treatment. These concepts are analogously defined for exported varieties using the superscript X as opposed to M.

Using these notations, equation (3) can be written as

$$EV_t = \underbrace{\sum_{g \in G^M} R_{g0} \cdot ATT_{gt}^{qm}}_{\equiv QM_t} + \underbrace{\left(- \sum_{g \in G^M} M_{g0} \cdot ATT_{gt}^{pm*} \right)}_{\equiv PM_t} + \underbrace{\sum_{g \in G^X} X_{g0} \cdot ATT_{gt}^{px*}}_{\equiv PX_t}, \tag{4}$$

where $R_{g0} = \sum_{ij \in \Omega_g^M} r_{ij0}$, $M_{g0} = \sum_{ij \in \Omega_g^M} m_{ij0}$, and $X_{g0} = \sum_{ij \in \Omega_g^X} x_{ij0}$ are aggregated trade values specific to each treatment group. The ATT terms in the equation are given by

$$ATT_{gt}^{qm} = \sum_{ij \in \Omega_g^M} \left(\frac{r_{ij0}}{R_{g0}} \right) \Delta \ln q_{ijt}^m \tag{5}$$

$$ATT_{gt}^{pm*} = \sum_{ij \in \Omega_g^M} \left(\frac{m_{ij0}}{M_{g0}} \right) \Delta \ln p_{ijt}^{m*} \tag{6}$$

$$ATT_{gt}^{px*} = \sum_{ij \in \Omega_g^X} \left(\frac{x_{ij0}}{X_{g0}} \right) \Delta \ln p_{ijt}^{x*} \tag{7}$$

For instance, ATT_{gt}^{pm*} measures the average effect of tariff changes on the pre-tariff prices of imported varieties in group g, weighted by their initial import values.

If heterogeneity in ATTs across treatment groups is not of interest, we can instead express equation (4) in terms of aggregate ATTs as in

$$EV_t = R_0 \cdot ATT_t^{qm} + (-M_0 \cdot ATT_t^{pm*}) + X_0 \cdot ATT_t^{px*}, \tag{8}$$

where $R_0 = \sum_{g \in G^M} R_{g0}$, $M_0 = \sum_{g \in G^M} M_{g0}$, and $X_0 = \sum_{g \in G^X} X_{g0}$ are aggregated trade values of treated varieties. The aggregate ATT terms in the equation are given by

$$ATT_t^{qm} = \sum_{g \in G^M} \left(\frac{R_{g0}}{R_0} \right) ATT_{gt}^{qm} \tag{9}$$

$$ATT_t^{pm*} = \sum_{g \in G^M} \left(\frac{M_{g0}}{M_0} \right) ATT_{gt}^{pm*} \tag{10}$$

$$ATT_t^{px*} = \sum_{g \in G^X} \left(\frac{X_{g0}}{X_0} \right) ATT_{gt}^{px*} \tag{11}$$

Equation (4) or (8) provides clear guidance for the empirical work. We first need to estimate the ATTs in the equations, weighting varieties by their appropriate initial trade values. Then, the ATTs can be combined using these equations to calculate the equivalent variation.

3.2 Difference-in-Differences (DiD) Identification

A natural way to estimate the ATT terms from the previous subsection is to use a difference-in-differences (DiD) event-study design. For example, consider the following two-way fixed effects (TWFE) specification for logged pre-tariff import prices:

$$\ln p_{ijt}^{m*} = \alpha_{ij} + \delta_t + \sum_{k \neq -1} \beta_k \cdot 1\{t = T_{ij} + k\} + \epsilon_{ijt}, \tag{12}$$

where α_{ij} is a variety fixed effect, δ_t is a time fixed effect, and $1\{t = T_{ij} + k\}$ indicates that month t is k months relative to the treatment date T_{ij} for variety ij. The month immediately before treatment is omitted as the reference period. The coefficient $\hat{\beta}_k$ is then typically interpreted as the average difference in outcomes between treated and untreated varieties at relative time k, compared with the month prior to treatment. Consistent with the welfare expressions in Section 3.1, observations are weighted by their 2017 import values.

To interpret $\hat{\beta}_k$ causally as an ATT, three assumptions are required. First, the no-anticipation assumption requires exporters not to change their prices in response to future tariffs before treatment occurs. Second, the parallel-trends assumption requires the treated-control difference in the reference pre-treatment period to be a valid counterfactual for the

treated-control difference in post-treatment periods had there been no treatment. Third, the stable unit treatment value assumption (SUTVA) requires the potential outcome of a variety to depend only on its own treatment status and not on the treatment status of other varieties.

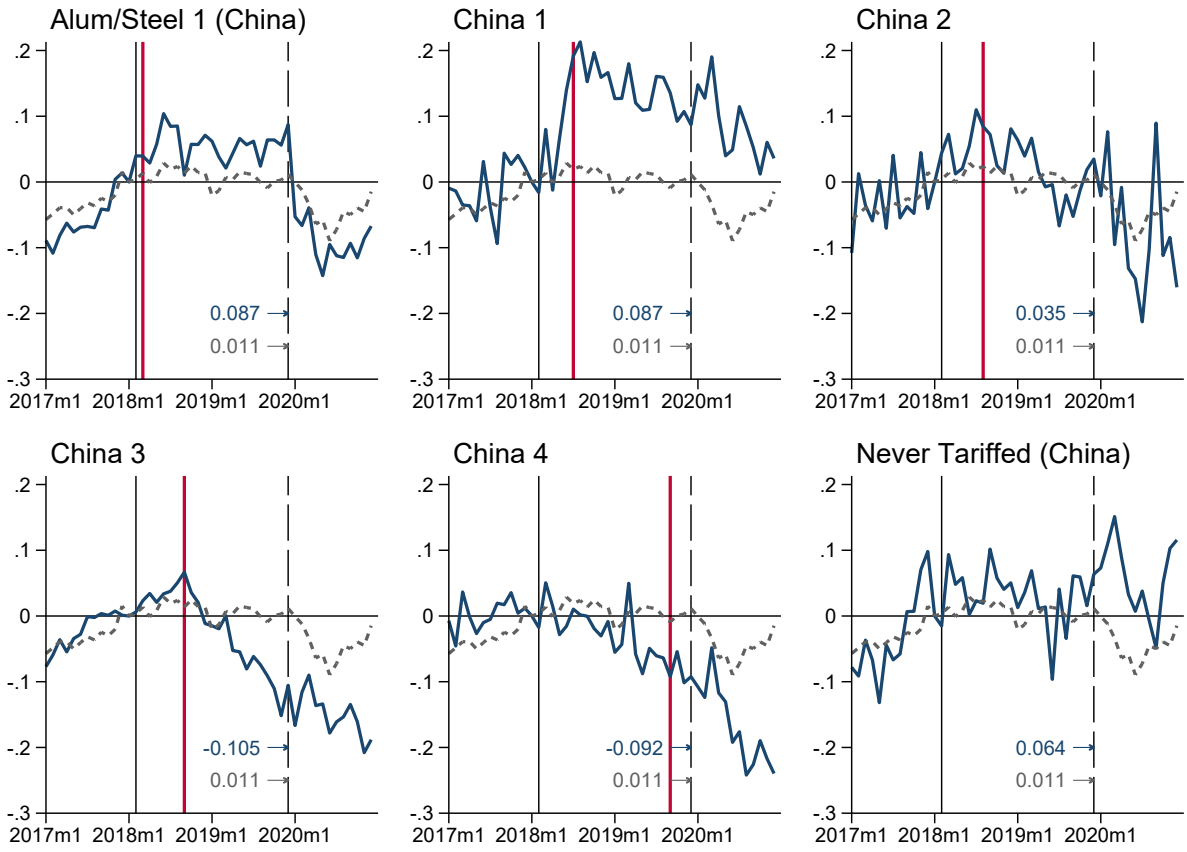
In this setting, the key identification choices are the definition of the control group and the timing of treatment. Previous studies have typically defined the control group to include both never-tariffed and not-yet-tariffed observations, while using tariff implementation dates as treatment dates. For the 2018–2019 trade war, both choices are difficult to defend a priori.

The standard control-group definition is problematic because some observations that were never ultimately tariffed were nonetheless affected by related trade policies or credible tariff threats. One example is steel imports from countries like Australia that avoided tariffs through bilateral trade deals but were instead subject to quotas. More importantly, a vast majority of Chinese varieties that were never tariffed during the sample period were publicly listed for tariffs scheduled to take effect on December 15, 2019. These product lists were announced by the U.S. on May 17, 2019 and by China on August 23, 2019, but the tariffs were later suspended following the Phase One agreement at the end of 2019. Treating these Chinese varieties as untreated is therefore difficult to justify, since their outcomes may already reflect direct exposure to announced tariff threats.

Using not-yet-tariffed observations as controls and using implementation dates as treatment dates is also problematic because tariffs were often announced well in advance, frequently with explicit product lists. For example, the fourth tariff wave against China was publicly announced on May 17, 2019 but implemented only on September 1, 2019 (Federal Register, Vol. 84, No. 96). Chinese exporters therefore had more than three months to adjust prices before implementation. More broadly, by late 2018 there were already public signals that the U.S. was considering tariffs on all remaining imports from China. For instance, President Trump communicated about levying tariffs on all remaining imports from China following U.S. and China's third exchange of tariffs (China 3) during an interview on September 7, 2018 (Rappeport, 2018). It was later followed by an official White House statement on September 17, 2018 (The White House, 2018). These details suggest that the no-anticipation assumption is unlikely to hold if tariff implementation dates are used as treatment dates.

Figure 2 provides ex-post evidence consistent with these concerns. It plots weighted-average changes in log import prices for Chinese products relative to products from other countries that were not subject to additional U.S. tariffs during 2018–2019. Two patterns are especially informative. First, prices for the China 4 group begin to decline well before the September 2019 implementation date. This pattern is difficult to reconcile with a

Figure 2: Relative Import Price Changes of Products from China and ROW



Notes: Each panel compares the weighted-average change in log import prices for products from China (solid line) against a comparison group of products from other countries that were not subject to additional U.S. tariffs in 2018–2019 (dashed line). Variety-level (10-digit product code by country) log price changes are calculated relative to Jan. 2018 and aggregated using 2017 annual import values as weights. The solid red vertical line denotes the month tariffs were first implemented for the given treatment group. The solid and dashed black vertical lines mark the beginning (Feb. 2018) and end (Dec. 2019) of the analyzed trade war period, respectively. The values near the arrows indicate the cumulative price change as of Dec. 2019.

no-anticipation assumption defined around tariff implementation and implies that not-yet-tariffed observations are poor controls. In particular, if China 4 prices were already falling before September 2019, then using those observations as controls for earlier-treated Chinese products will positively bias estimated treatment effects upward toward zero for other Chinese varieties.

Second, the never-tariffed Chinese products do not behave like an obviously clean untreated group. Their price path is broadly similar to that of never-tariffed products from the rest of the world through much of the sample, which alleviates the concern that they experienced large anticipatory price declines. However, they also exhibit somewhat higher price growth during 2019. By December 2019, never-tariffed Chinese products show a price increase of about 6.4%, compared with about 1.1% for never-tariffed products from the rest

of the world. This pattern suggests that including never-tariffed Chinese observations in the control group may introduce negative bias.

Taken together, these considerations motivate two design choices for the analysis going forward. First, for identification purposes, I treat the beginning of the trade war in February 2018 as the relevant treatment date rather than using tariff implementation dates. Second, my preferred specifications exclude observations that are especially likely to violate SUTVA from the control group. These are never-tariffed aluminum/steel and Chinese varieties. The next section shows how these choices affect the TWFE estimates.

3.3 Preliminary Evidence from TWFE Estimates

Figure 3 reports TWFE estimates from equation (12), using February 2018 as the common treatment date for all tariffed varieties and January 2018 as the omitted reference period. The purpose of this exercise is not to present the final estimator, but to show how the estimated price effects vary across different fixed-effects specifications once treatment timing is defined in a way that is more consistent with the identification discussion in Section 3.2. Taken together, these results show that once treatment timing is aligned with the start of the trade war, conventional TWFE specifications reveal sizable declines in pre-tariff import prices.

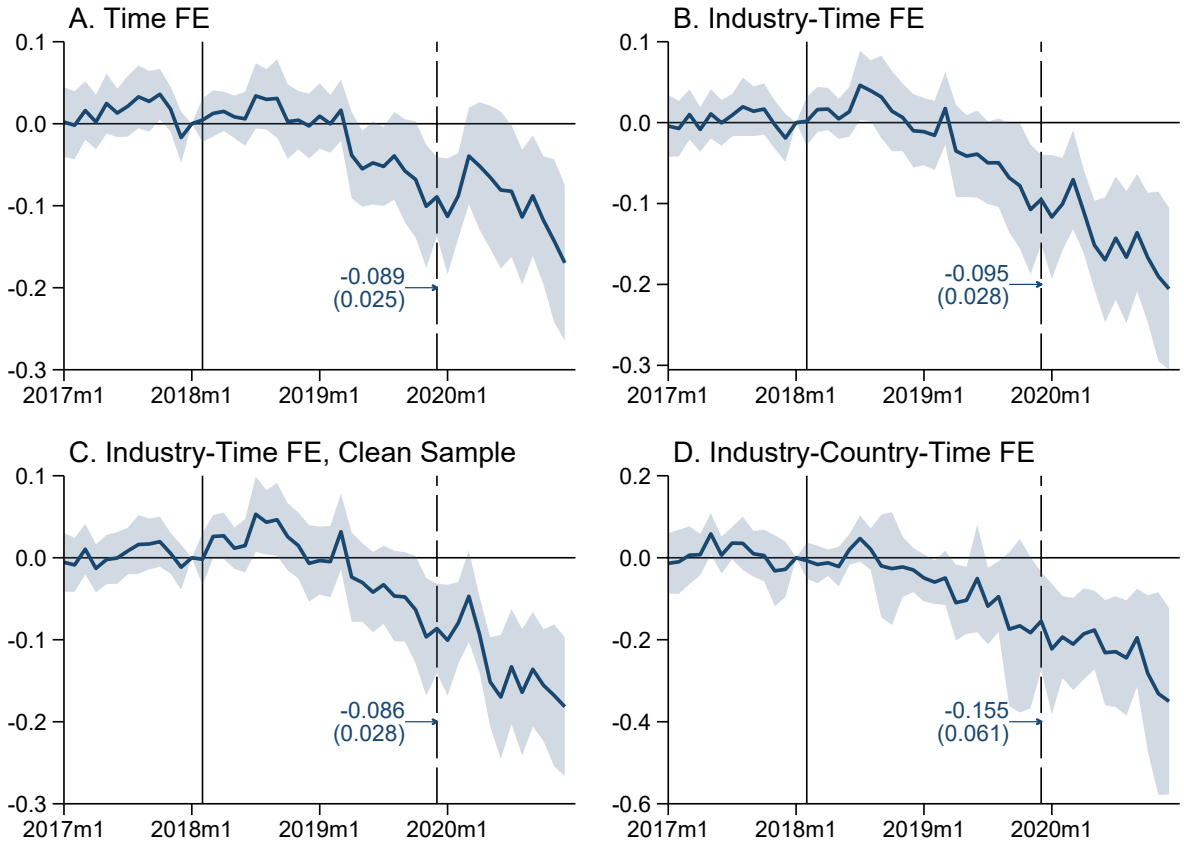
Panel A presents the basic specification with variety and time fixed effects. Pre-tariff import prices are roughly flat through much of 2018 and then begin to decline in 2019. By December 2019, I estimate a statistically significant and economically meaningful decline. However, the pre-treatment coefficients deviate somewhat from zero.

Panel B replaces the time fixed effects with industry-time fixed effects, where industries are defined at the 3-digit NAICS level. This specification compares treated and untreated varieties within the same industry in each period and therefore makes the parallel-trends assumption more plausible. The estimated decline by December 2019 is similar to that in Panel A, but the pre-trends are tighter and the post-2018 decline is more consistent.

Panel C applies the same specification as Panel B to a cleaner estimation sample that excludes never-tariffed aluminum/steel and Chinese varieties. The resulting dynamics in the estimates remain similar to Panel B. The December 2019 effect is slightly smaller in magnitude, which is consistent with the idea that including these observations in the control group may induce a modest negative bias.

For completeness, Panel D adds industry-country-time fixed effects. This specification is attractive in its ability to absorb time-varying trends at the country level (e.g., exchange-rate movements). However, for Chinese varieties it necessarily compares treated Chinese varieties

Figure 3: TWFE Estimates of Effects on Pre-tariff Import Prices



Notes: Panel A plots the estimates from the TWFE specification in equation (12). Panel B and C plot the estimates from alternative TWFE specifications that include industry-time fixed effects but Panel C further excludes never-tariffed aluminum/steel and Chinese varieties from the sample. Panel D plots the estimates from a TWFE specification with industry-country-time fixed effects. Observations are weighted by their 2017 import values. Shaded regions are 95% confidence intervals based on standard errors clustered at the 8-digit product level. The solid and dashed black vertical lines mark the beginning (Feb. 2018) and end (Dec. 2019) of the analyzed trade war period, respectively. The point estimates (standard errors in parentheses) for Dec. 2019 are displayed near the arrows.

to untreated Chinese varieties within the same industry and period. If those untreated Chinese varieties are themselves contaminated by tariff threats, this comparison is not a valid control and so it is not preferred. Indeed, the larger estimated decline in Panel D is consistent with the pattern in Figure 2 that never-tariffed Chinese varieties experienced relatively higher price increase than never-tariffed varieties from other countries.

3.4 Comparison with Previous Approaches

The evidence so far suggests that pre-tariff import prices did decline during the 2018–2019 trade war. Why, then, did earlier studies typically fail to detect such declines? A key reason

is that much of the literature has estimated variants of the following three-way fixed effects (3WFE) specification:

$$\ln p_{ijt}^{m*} = \alpha_{ij} + \lambda_{it} + \eta_{jt} + \sum_{k \neq -1} \beta_k \cdot 1\{t = T_{ij} + k\} + \epsilon_{ijt}, \tag{13}$$

where α_{ij} is a variety fixed effect, λ_{it} is a product-time fixed effect, and η_{jt} is a country-time fixed effect. This specification is attractive because it absorbs many time-varying confounders, such as product-specific demand shocks and country-specific exchange-rate movements.

However, $\hat{\beta}_k$ in this specification should not be interpreted as a standard DiD estimator but as a triple-difference or difference-in-difference-in-differences (DDD) estimator instead. To build intuition, consider a simplified two-period setting in which treatment applies only to Chinese varieties and all non-Chinese varieties are pooled into a single comparison group. In this case, Olden and Møen (2022) show that the 3WFE specification in equation (13) is equivalent to a standard triple-difference setup:

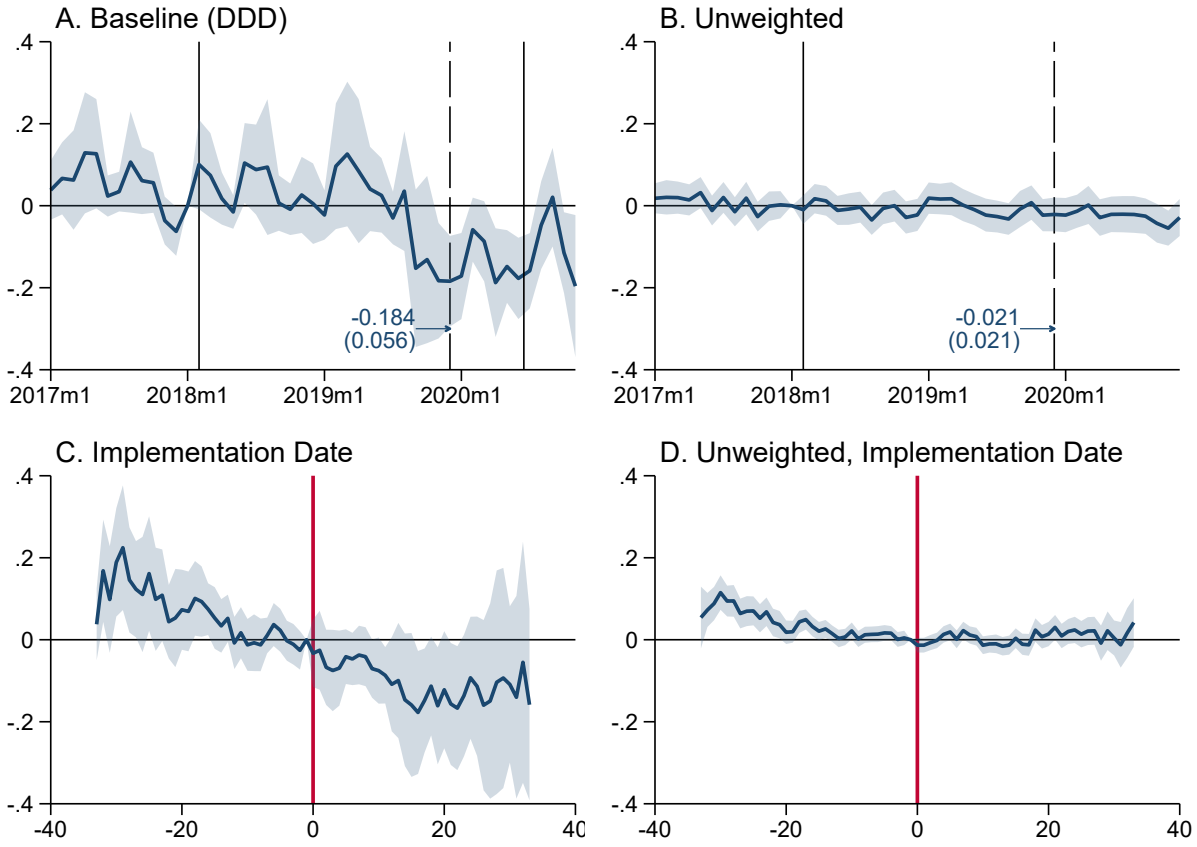
$$\begin{aligned} \ln p_{ijt}^{m*} = {} & \beta_0 + \beta_1 \text{Treat}_i \times \text{China}_j \times \text{Post}_t \\ & + \beta_2 \text{Treat}_i \times \text{China}_j + \beta_3 \text{Treat}_i \times \text{Post}_t + \beta_4 \text{China}_j \times \text{Post}_t \\ & + \beta_5 \text{Treat}_i + \beta_6 \text{China}_j + \beta_7 \text{Post}_t + \epsilon_{ijt}, \end{aligned}$$

and the coefficient on the triple interaction, β_1, is the DDD estimand. Ortiz-Villavicencio and Sant'Anna (2025) discuss the 3WFE specification in a more general setting.

As Olden and Møen (2022) show, the coefficient of interest, β_1, is a triple-difference estimator. Intuitively, it can be understood as the difference between two DiD estimators that compare the mean difference in outcomes between treated and untreated products for China and for other countries, respectively. The third difference then compares these two DiD estimates. This can be useful because DDD identification does not require parallel trends to hold separately within China and the rest of the world. Instead, it requires any deviation from parallel trends to evolve similarly across those two groups.

This distinction matters for two reasons. First, the identifying assumption for the DDD specification is different from the parallel-trends assumption used in Section 3.2. Either one may hold without the other. Second, the DDD design requires untreated observations within each comparison group. If the view is that no Chinese varieties were untreated during this trade war, then the DDD interpretation becomes difficult to sustain. These concerns become even more acute when treatment timing is staggered and implementation dates are used as treatment dates.

Figure 4: 3WFE Estimates of Effects on Pre-tariff Import Prices



Notes: Panel A plots the estimates from the 3WFE specification in equation (13). Panel B plots the estimates from the same specification as in Panel A but without weighting observations. Panel C plots the estimates from the 3WFE specification in equation (13) using tariff implementation dates as the treatment dates. Panel D plots the estimates from the same specification as in Panel C but without weighting observations. Observations are weighted by their 2017 import values in Panels A and C. Shaded regions are 95% confidence intervals based on standard errors clustered at the 8-digit product level. The solid and dashed black vertical lines mark the beginning (Feb. 2018) and end (Dec. 2019) of the analyzed trade war period, respectively. The point estimates (standard errors in parentheses) for Dec. 2019 are displayed near the arrows.

Figure 4 shows how these choices matter in practice. Panel A uses February 2018 as the common treatment date and weights observations by their 2017 import values. Under this specification, pre-tariff import prices do decline, but the estimated decline appears later and both pre- and post-treatment coefficients are much noisier than in the TWFE specifications. Panel B keeps the same treatment timing but does not use weights. The estimated price response then becomes close to zero throughout the sample, which reproduces the broad conclusion of the earlier literature.

Panels C and D instead use tariff implementation dates as treatment dates. Panel C retains import-value weights and still shows falling pre-tariff prices, but it also displays a clear pattern of declines prior to implementation. Panel D, which is closest to the specifications

used in previous studies, combines implementation-date timing with unweighted estimation. It shows little evidence of any post-treatment price decline while still showing statistically significant pre-trends. This pre-trend may have gone unnoticed in previous studies, because they typically trim event-study windows (e.g., 12 months before the treatment date).

These results indicate that weighting observations is the main factor driving the difference between my findings and previous results in the literature.

3.5 Main Empirical Specification

The TWFE results above are useful, but they are not ideal for the remainder of the paper. In particular, the empirical analysis that follows requires an estimator that can flexibly accommodate heterogeneous treatment effects across tariff waves. Thus, I adopt the imputation-based estimator of Borusyak et al. (2024) as the main estimator for the rest of the paper.

Their estimator can be described in three steps. First, I use untreated observations only, excluding the observations identified in Section 3.2 as potentially problematic controls, to estimate untreated outcome dynamics:

$$\ln p_{ijt}^{m*} = \alpha_{ij}^{NT} + \delta_{st}^{NT} + \varepsilon_{ijt}^{NT}, \tag{14}$$

where δ_{st}^{NT} denotes industry-time fixed effects. Relative to the TWFE specifications above, the key difference is that these fixed effects are estimated using untreated observations only. Second, I impute untreated potential outcomes for treated observations using the estimated fixed effects and recover observation-level treatment effects as

$$\widehat{\theta}_{ijt}^{pm*} = \ln p_{ijt}^{m*} - \left(\widehat{\alpha}_{ij}^{NT} + \widehat{\delta}_{st}^{NT}\right). \tag{15}$$

Thus, $\widehat{\alpha}_{ij}^{NT} + \widehat{\delta}_{st}^{NT}$ serves as the estimated untreated outcome for variety ij in period t. Third, I aggregate these observation-level effects using initial trade values as weights to obtain either aggregate or group-specific ATTs. For aggregate effects, I use

$$\widehat{ATT}_t^{pm*} = \sum_{g \in G^M} \sum_{ij \in \Omega_g^M} \left(\frac{m_{ij0}}{M_0}\right) \widehat{\theta}_{ijt}^{pm*}, \tag{16}$$

and for treatment-group-specific effects,

$$\widehat{ATT}_{gt}^{pm*} = \sum_{ij \in \Omega_g^M} \left(\frac{m_{ij0}}{M_{g0}}\right) \widehat{\theta}_{ijt}^{pm*}. \tag{17}$$

Other aggregations can be considered in an analogous manner.

Under no-anticipation and parallel trends assumptions, Borusyak et al. (2024) show that these estimators are unbiased. In the present setting, the parallel-trends restriction is that untreated outcomes can be decomposed into variety fixed effects and industry-time fixed effects. Because this restriction is imposed on untreated outcomes for all periods, it can be tested directly using untreated observations. To do so, I estimate a pre-trends regression that interacts treatment-group indicators with leads relative to the common treatment date:

$$\ln p_{ijt}^{m*} = \alpha_{ij}^{pre} + \delta_{st}^{pre} + \sum_{g \in G^M} \sum_{k<-1} \beta_{gk}^{pre} \cdot 1\{G_{ij} = g\} \cdot 1\{t = T_{ij} + k\} + \varepsilon_{ijt}^{pre}. \tag{18}$$

Then, I test whether the coefficients $\hat{\beta}_{gk}^{pre}$ are different from zero. When testing for aggregate pre-trends, the group indicator variables are combined into a single indicator of treatment.

The imputation-based estimator also allows for a straightforward construction of average pass-through rates (APTR) of import tariffs. For each treated observation, I divide the estimated change in the pre-tariff price by the corresponding tariff change then add one:

$$1 + \frac{\widehat{\theta}_{ijt}^{pm*}}{\Delta \ln(1 + \tau_{ijt}^{m})}. \tag{19}$$

If, for example, there was no price response to the import tariff (i.e., $\widehat{\theta}_{ijt}^{pm*} = 0$), then the pass-through rate will be 1 (complete). If instead, foreign exporters lowered their price in response (i.e., $\widehat{\theta}_{ijt}^{pm*} < 0$), then the pass-through rate will be less than 1 (incomplete). These estimates can then be aggregated in a manner analogous to the aggregation of $\widehat{\theta}_{ijt}^{pm*}$ to obtain APTR estimates. Appendix Section B.1 details how the estimation and aggregation of tariff pass-through rates can be facilitated using the `did_imputation` Stata package of Borusyak et al. (2024).

A practical limitation of this estimator is that valid inference requires averaging across a sufficient number of observations. When observations are weighted, Borusyak et al. (2024) recommend at least 30 effective observations.[5] This threshold is easily satisfied for aggregate effects, but not always for narrowly defined treatment groups, especially on the export side where retaliatory tariffs were applied to much narrower sets of products. Accordingly, I suppress standard errors for treatment groups whose effective number of observations falls below this threshold on average and report only their point estimates for completeness.[6]

[5] The effective number of observations is calculated as the inverse Herfindahl index using the weights as the shares.

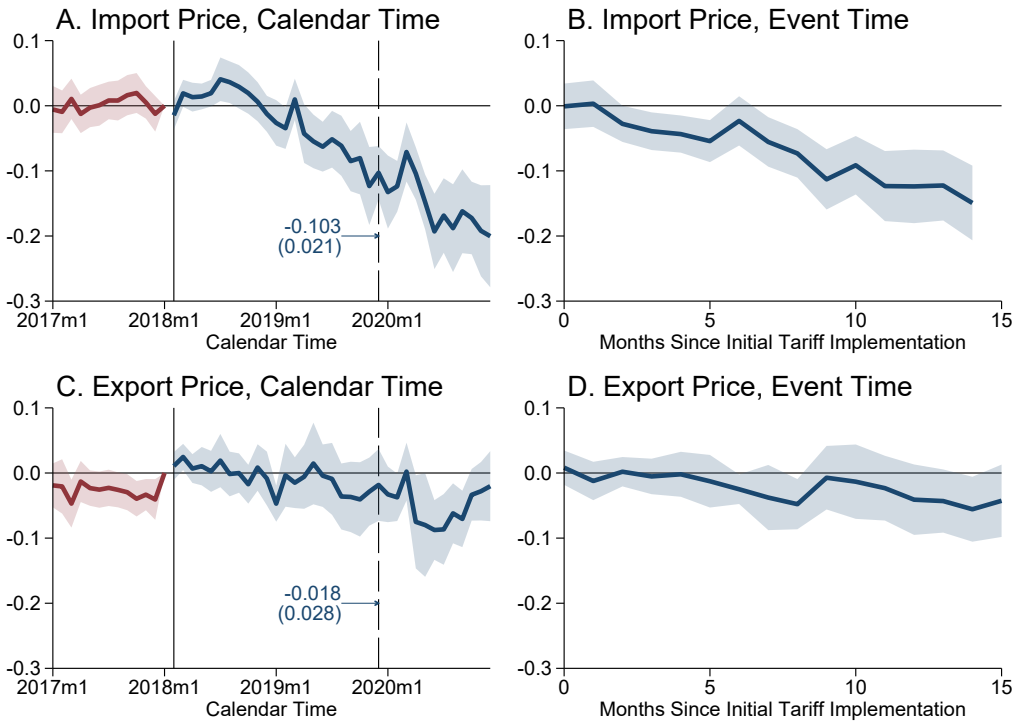
[6] When looking at import prices, these groups are Solar/Washer, Aluminum (KOR), and E.U. Dispute. For export prices, all but two groups Aluminum/Steel: Canada and China 3 have insufficient effective observations.

4 Effects on U.S. Prices and Welfare

4.1 Aggregate Effects on Pre-Tariff Prices

Figure 5 reports the main estimated effects of the 2018–2019 trade war tariffs on U.S. pre-tariff import and export prices using the imputation-based estimator described in Section 3.5. Panel A and C present the aggregate ATT estimates in calendar time, while Panels B and D present the same estimates reorganized in event time. Details on reorganizing the estimates in event time are provided in Appendix Section B.2.

Figure 5: Estimated Aggregate Effects on Pre-Tariff Import and Export Prices



Notes: Panel A and C plot in blue the ATT estimates by calendar time in equation (16) for pre-tariff import and export prices, respectively. The pre-trend estimates are in red. Panel B and D plot the ATT estimates by event time (months since tariff implementation) from equation (21) instead. I report only event times for which data from all treatment groups are available. Shaded regions are 95% confidence intervals based on standard errors clustered at the 8-digit product level (6-digit for export prices). The solid and dashed black vertical lines mark the beginning (Feb. 2018) and end (Dec. 2019) of the analyzed trade war period, respectively. The point estimates (standard errors in parentheses) for Dec. 2019 are displayed near the arrows.

Panel A shows a clear decline in pre-tariff import prices. Prices rise modestly in the earliest months of the trade war but then decline steadily beginning in mid-2018. By December 2019, the estimated decline is 10.3% and is statistically significant. With the brief exception of March 2020, this decline persists through 2020. The aggregated pre-trend estimates remain close to zero throughout 2017, which supports the identifying assumptions. Panel

B presents the same effects on import prices in event time, where month zero corresponds to each variety's initial tariff implementation date. The estimates show a steady decline following implementation, reaching 14.9% after 14 months.

In contrast, Panel C and D show much weaker evidence of declines in pre-tariff export prices. In calendar time, the estimated decline by December 2019 is only 1.8% and is not statistically significant. Although the estimates become more negative in 2020, this appears more consistent with a break around the pandemic period than with a continuation of the 2018–2019 pattern. The corresponding pre-trend estimates are also less convincing than on the import side. Although mostly flat, there is a noticeable jump in January 2018. The magnitude of the coefficients suggests that using an alternative reference period will do little to impact the results in a meaningful way. Estimates presented in event time lead to the same conclusion.

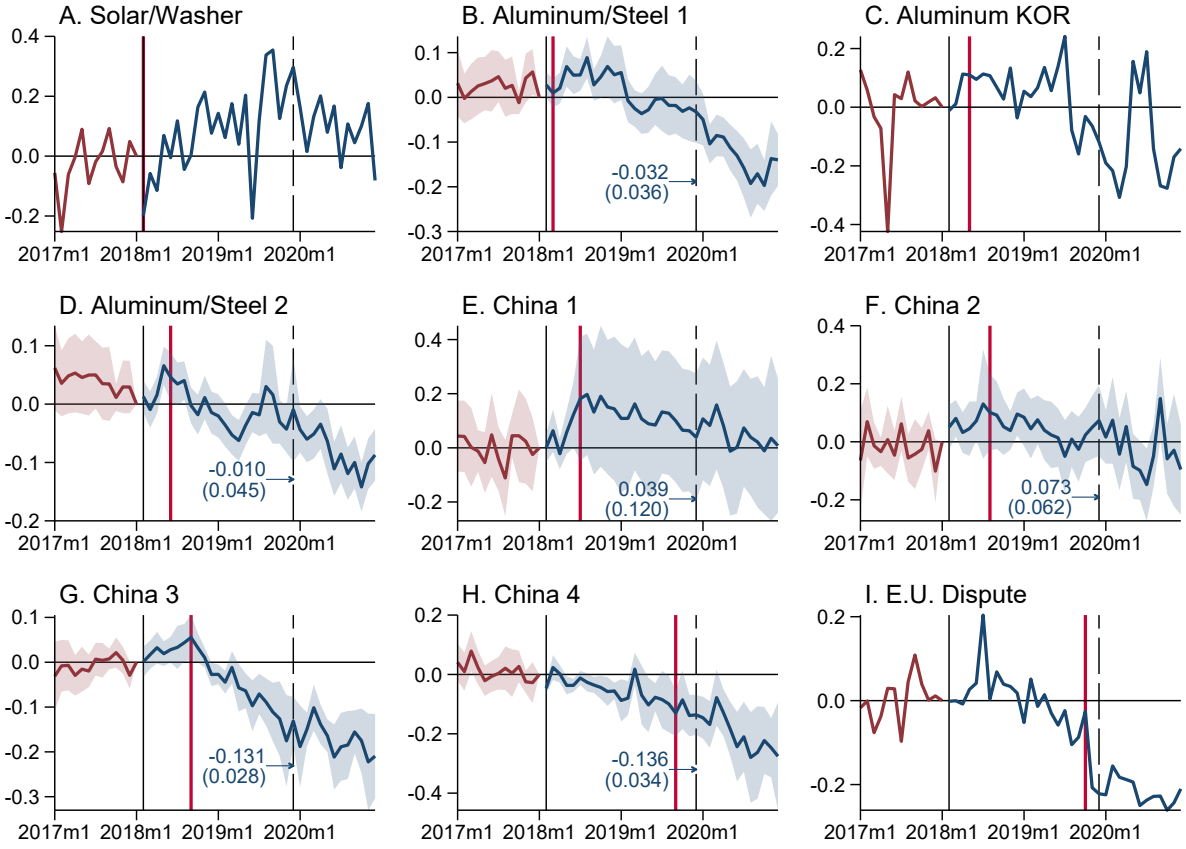
Overall, Figure 5 provides strong evidence that foreign exporters lowered pre-tariff prices in response to U.S. tariffs, but much weaker evidence that U.S. exporters lowered prices in response to retaliatory tariffs.

4.2 Effects on Pre-Tariff Prices by Treatment Group

The aggregate estimates in Figure 5 mask substantial heterogeneity across treatment groups. Figure 6 reports the ATT estimates for each of the nine treatment groups on the import side. By December 2019, statistically significant declines in pre-tariff import prices appear only for China 3 and China 4. The estimated declines are 13.1% and 13.6%, respectively. For Aluminum/Steel 1-2 and China 1-2, I fail to reject the null of no price change by the end of 2019. Standard errors are not reported for Solar/Washer, Aluminum (KOR), and the E.U. Dispute groups because their effective number of observations is too small for reliable inference, although I report the point estimates for completeness.

Although only two treatment groups display statistically significant import-price declines, China 3 and China 4 account for the lion's share of import value exposed to U.S. tariffs during the trade war. Their combined import value was roughly $300 billion in 2017 or 73% of all tariffed imports, so their price responses dominate the aggregate ATT on import prices. The patterns for these two groups are also the most convincing visually. In both panels, the pre-trend estimates remain close to zero and do not suggest that the post-2018 declines are simply picking up pre-existing trends. The estimated treatment effects then become steadily more negative after the beginning of the trade war and remain negative through 2020. In the case of China 4, the estimates also reinforce the earlier evidence that prices began to decline well before the September 2019 implementation date, consistent with anticipatory

Figure 6: Estimated Effects on Pre-tariff Import Prices by Treatment Group

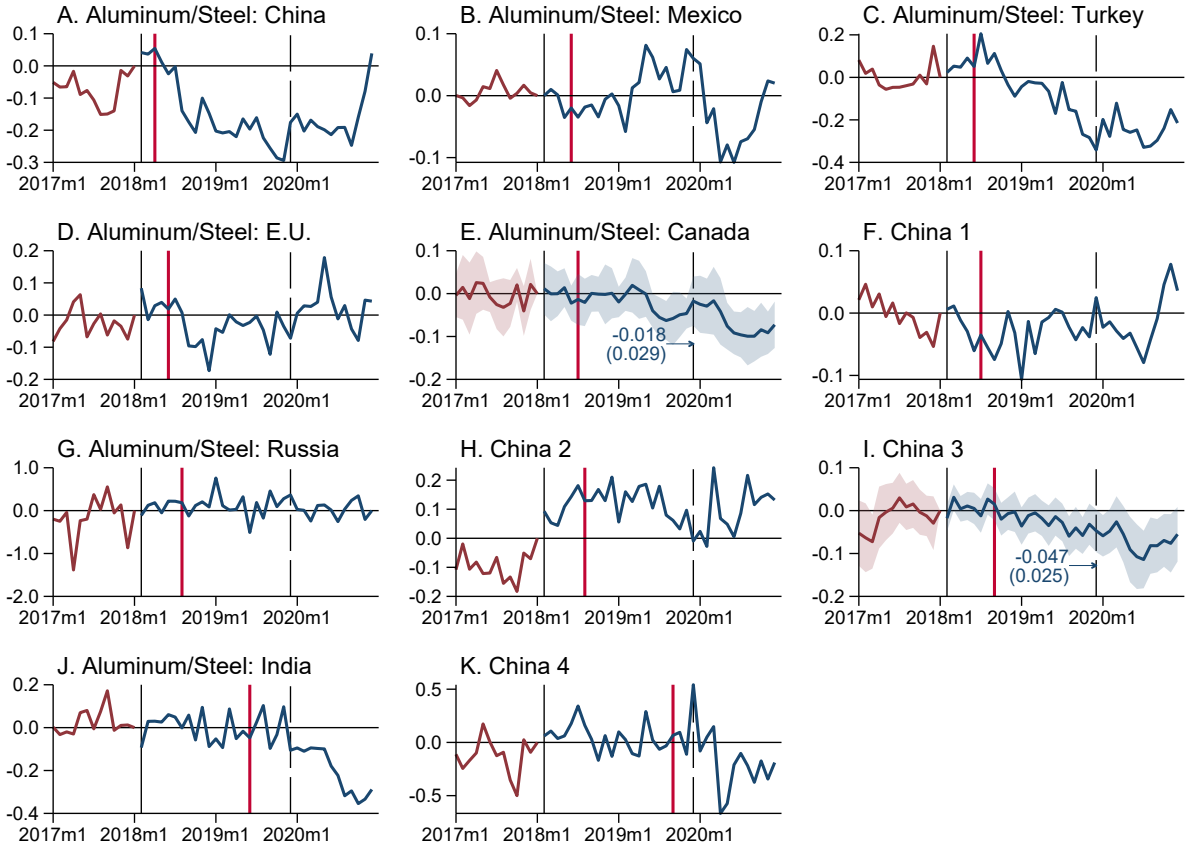


Notes: Each panel plots the Borusyak et al. (2024) estimates in equation (17) in blue and the pre-trend estimates from equation (18) in red. Shaded regions are 95% confidence intervals based on standard errors clustered at the 8-digit product level. The solid red vertical line denotes the month tariffs were first implemented for the given treatment group. The solid and dashed black vertical lines mark the beginning (Feb. 2018) and end (Dec. 2019) of the analyzed trade war period, respectively. The point estimates (standard errors in parentheses) for Dec. 2019 are displayed near the arrows.

responses to tariff announcements.

Figure 7 reports the corresponding estimates for the eleven treatment groups on the export side. Here the effective-sample-size problem is much more severe. Because retaliatory tariffs were applied to narrower product sets than the U.S. tariffs, the weighted estimates within many export treatment groups are dominated by a small number of observations, making standard errors unreliable. I therefore present these estimates mainly for transparency and limit interpretation to the two groups for which inference is more reliable: Aluminum/Steel: Canada (Panel E) and China 3 (Panel I). For Canada's retaliation, the estimated dynamics are mostly flat, dipping slightly toward the end of 2019. Nevertheless, the estimate for December 2019 is not statistically different from zero. For China 3, I find a statistically significant decline of 6.6% in export prices by the end of 2019. However, the

Figure 7: Estimated Effects on Pre-tariff Export Prices by Treatment Group



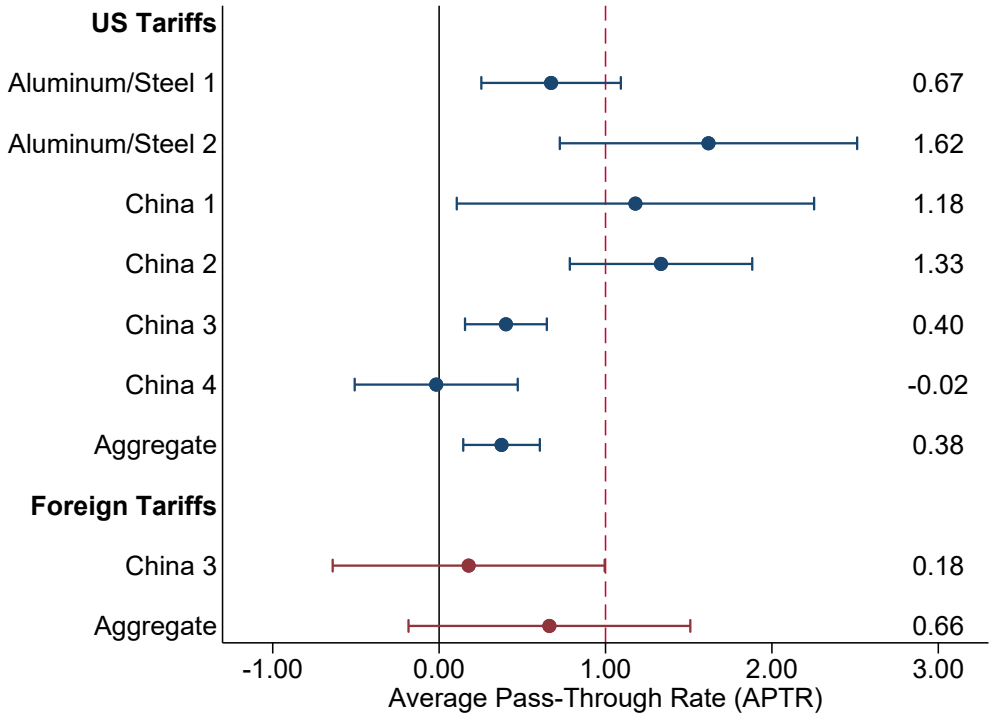
Notes: Each panel plots the Borusyak et al. (2024) estimates in equation (17) in blue and the pre-trend estimates from equation (18) in red. Shaded regions are 95% confidence intervals based on standard errors clustered at the 8-digit product level. The solid red vertical line denotes the month tariffs were first implemented for the given treatment group. The solid and dashed black vertical lines mark the beginning (Feb. 2018) and end (Dec. 2019) of the analyzed trade war period, respectively. The point estimates (standard errors in parentheses) for Dec. 2019 are displayed near the arrows.

pre-trends are worrisome. Despite a flatter trend leading up to the trade war, there are signs of statistically significant deviations earlier on in 2017.

4.3 Tariff Pass-Through Rates

I now turn to the average pass-through rates of U.S. and foreign tariffs. These estimates are central because they summarize how the tariff burden was split between importers and exporters. An APTR of one corresponds to complete pass-through, meaning that exporters did not lower their pre-tariff prices. An APTR below one implies incomplete pass-through, meaning that exporters absorbed part of the tariff by reducing their prices. By the end of 2019, there is strong evidence that the pass-through of U.S. tariffs was incomplete, but much weaker evidence on the pass-through of retaliatory tariffs.

Figure 8: Average Pass-Through Rates of U.S. and Foreign Import Tariffs



Notes: This figure plots the estimated average pass-through rate (APTR) of U.S. and foreign import tariffs for each treatment group and in aggregate for December 2019. Each estimate is obtained by aggregating relevant observation-level pass-through estimates in equation (19). The point estimates for the pass-through rates are listed in the right margin. Results for treatment groups with insufficient effective number of observations are not shown. The capped bars are 95% confidence intervals based on standard errors clustered at the 8-digit product level for imports and at the 6-digit level for exports.

Figure 8 reports the estimated APTRs for December 2019. On the import side, the aggregate APTR for U.S. tariffs is 38% and precisely estimated. Importantly, the null of a complete pass-through can be rejected decisively. The estimate implies that foreign exporters absorbed the majority of U.S. tariff burden by lowering their prices. Consistent with the heterogeneity documented in Section 4.2, APTRs also vary considerably across treatment groups. A complete pass-through rate can be rejected at the 5% level only for China 3 and China 4, the two groups that displayed largest price declines previously. For China 4, I fail to reject the null that foreign exporters fully absorbed the tariff increases (i.e., zero tariff pass-through). This evidence contrasts with earlier work that concluded that the pass-through of U.S. tariffs remained nearly complete through 2019 (Amiti et al., 2020).

On the export side, the evidence is considerably weaker. The only group-specific retaliatory-tariff APTR reported in Figure 8 is for China 3, and the aggregate APTR is 66% but imprecisely estimated.[7] In particular, I cannot reject the null that the pass-through was complete.

[7]The APTR for Canada's retaliatory tariffs cannot be computed for December 2019 as they were lifted

However, these estimates should be interpreted with caution because the effects on export prices in Figure 5 are imprecisely estimated.

4.4 Welfare

I now return to the equivalent variation in equation (8) to assess the overall welfare effect of the 2018–2019 trade war tariffs on the U.S. economy. The second and third terms in that expression, corresponding to the effects on pre-tariff import and export prices, have already been estimated and presented in Figure 5. To estimate the first term, I use logged import quantity as the outcome variable and 2017 tariff revenues as weights. By December 2019, the tariff-revenue-weighted average decline in import quantities is 44.1%.

Table 2: Impact of the 2018–2019 Trade War on U.S. Welfare

	Welfare Effect (EV_t)	Import Quantity (QM_t)	Import Price (PM_t)	Export Price (PX_t)
% of GDP	0.17	−0.03	0.20	−0.01
Billion USD	32.95	−5.03	40.14	−2.16

Notes: Percentage of GDP calculated using U.S. GDP in 2017. Calculations are based on equation (8) and the corresponding ATT estimates for December 2019.

The three ATT terms in equation (8) are then multiplied by their respective initial trade values: $11.407 billion of tariff revenue, $390.339 billion of imports, and $118.095 billion of exports. The resulting welfare impacts for these terms and in total are summarized in Table 2. I estimate a 0.17% (or $32.95 billion) U.S. welfare *gain* at the end of 2019. The individual terms further show that this was driven largely by a net positive terms-of-trade gain worth about $38 billion. The tariff revenue effect (QM_t) from decreased import quantities is smaller, because the U.S. collected only small sums of tariff revenue in 2017. This welfare gain is considerably larger in magnitude than the welfare losses estimated by Amiti et al. (2019) and Fajgelbaum et al. (2020) for 2018 ($8.2 billion and $7.2 billion, respectively) and has the opposite sign.

This welfare gain should be understood as the short-run effect. It does not account for longer-run adjustments such as changes in investment, sourcing patterns, or supply-chain reorganization (Fajgelbaum et al., 2024; Alfaro and Chor, 2023; Freund et al., 2024; Garred and Yuan, 2025; Ma et al., 2025). For this reason, it is not directly comparable to studies that focus on general-equilibrium or long-run welfare effects, which often find negative overall

in May 2019.

consequences of the trade war (Fajgelbaum et al., 2020; Amiti et al., 2021; Lashkaripour, 2021).

5 Importer Market Power and Tariff Pass-Through

I close the paper by examining whether heterogeneous pass-through rates can be explained by bilateral importer market power. Standard trade theory predicts that tariffs imposed by importers with greater market power should induce larger reductions in exporters' pre-tariff prices. As argued in Section 2.2, overall importer market shares may not capture this relationship well in the context of the 2018–2019 trade war, since it heavily affected bilateral trade between the U.S. and China. I therefore study whether pass-through rates were systematically lower in products where the importer had greater bilateral market power over the exporter.

To implement this exercise, I partition treated import and export varieties into respective deciles of bilateral importer market power defined in equation (2). Within each decile d, I then calculate the average pass-through rate by aggregating the observation-level pass-through estimates from equation (19). For imports, I use

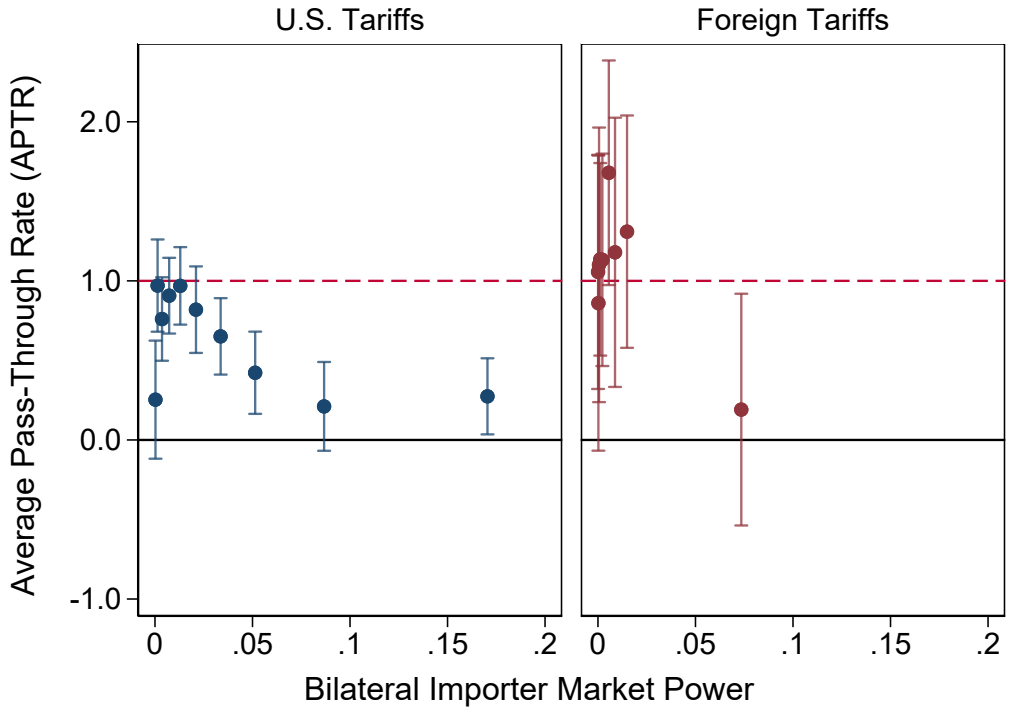
$$\widehat{APTR}_{dt}^{pm*} = \frac{1}{|\Omega_d^M|} \sum_{ij \in \Omega_d^M} \widehat{APTR}_{ijt}^{pm*}, \tag{20}$$

where Ω_d^M denotes the set of tariffed import varieties in the d-th decile of bilateral importer market power. The corresponding estimates for foreign retaliatory tariffs are constructed analogously. The decile-level estimates use unweighted averages here to isolate the role of importer market power. As a result, these estimates capture the pass-through rates for a typical variety in each decile.

Figure 9 plots the decile-level estimates for U.S. tariffs in the left panel and for foreign retaliatory tariffs in the right panel, together with 95% confidence intervals. The import-side pattern is clear. Pass-through rates are highest in the lower deciles of bilateral importer market power and decline steadily in the upper deciles. In particular, I fail to reject complete pass-through for most deciles in the lower half of the distribution. In contrast, statistically significant incomplete pass-through emerges in the upper tail, beginning around the seventh decile, where the average bilateral importer market power is about 3.4%. The estimated pass-through rates then continue to fall across the highest deciles, reaching values below 30%.

This pattern helps reconcile the earlier results in the paper. On the one hand, Section 4 showed that the aggregate pass-through of U.S. tariffs was incomplete. On the other hand,

Figure 9: Average Pass-Through Rates by Bilateral Importer Market Power Deciles



Notes: This figure plots average pass-through rates by deciles of bilateral importer market power, measured using 2017 CEPII-BACI data on 6-digit product level international trade flows. Each point aggregates observation-level pass-through estimates within a decile of bilateral importer market power. The estimate for the sixth decile is not shown for foreign tariffs due to a large 95% confidence interval $(-1.86, 5.70)$. The capped bars are 95% confidence intervals based on standard errors clustered at the 8-digit product level for imports and at the 6-digit level for exports.

Figure 9 shows that incomplete pass-through was not pervasive across all products. For more than half of treated import varieties, including the median variety, I cannot reject complete pass-through. Thus, the aggregate decline in import prices was driven disproportionately by products in which the U.S. had considerable bilateral market power over the exporter, rather than by a uniform price response across all tariffed imports.

This also helps explain the discrepancy between weighted and unweighted estimates for aggregate effects on import prices. Unweighted estimates would indicate nearly complete pass-through, because the pass-through was indeed complete for a typical import variety. Weighting estimates by their initial import values would naturally place greater weight on import varieties where the U.S. has higher market power over the exporter. This is because import value appears as a numerator in the bilateral importer market power. These varieties are the ones where the pass-through is incomplete, which is what the weighted estimate indicates.

The evidence for foreign retaliatory tariffs is also informative. The right panel of Figure 9

suggests that retaliating countries had very little bilateral importer market power over U.S. exporters for most treated export varieties. The average bilateral market power is only 1.5% even at the ninth decile, and for the first nine deciles I cannot reject complete pass-through. The top decile, where the average bilateral importer market power rises to 7.3%, is the lone exception with an estimated APTR of 19%. Admittedly, these estimates are considerably noisier than those for U.S. tariffs, and one decile is omitted from the figure because its confidence interval is extremely wide.

With this caveat in mind, I note that the magnitudes of these estimates are quite consistent with those for U.S. tariffs. For instance, a 1.5% bilateral market power corresponds to the fifth decile for the U.S. tariffs, where the pass-through is complete. Furthermore, a 7.3% bilateral market power falls closest to the ninth decile for the U.S. tariffs, which has a comparable pass-through rate of 21%. These results help explain the more subdued effects on export prices in Section 4. For the vast majority of treated export varieties, the foreign countries held no market power to induce lower prices in the first place.

6 Conclusion

This paper revisits the impact of the 2018–2019 trade war on U.S. prices and welfare. By December 2019, I estimate a 10.3% value-weighted decline in pre-tariff import prices and an aggregate pass-through rate of 38% for U.S. tariffs. In contrast, the evidence on export prices and on the pass-through of retaliatory tariffs is much weaker. These results overturn the view that U.S. tariffs were borne almost entirely by U.S. importers and consumers through 2019.

Combining the estimated effects, I find a U.S. welfare gain driven by a positive net terms-of-trade effect as standard trade theory would suggest for a large importer. The paper also shows that incomplete pass-through was concentrated in a smaller share of products where the importer had greater market power over the exporter. Overall, these results suggest that the terms-of-trade effects of tariffs depend not only on country size, but also on the bilateral importer market power.

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A Data Appendix

A.1 U.S. Import Tariffs

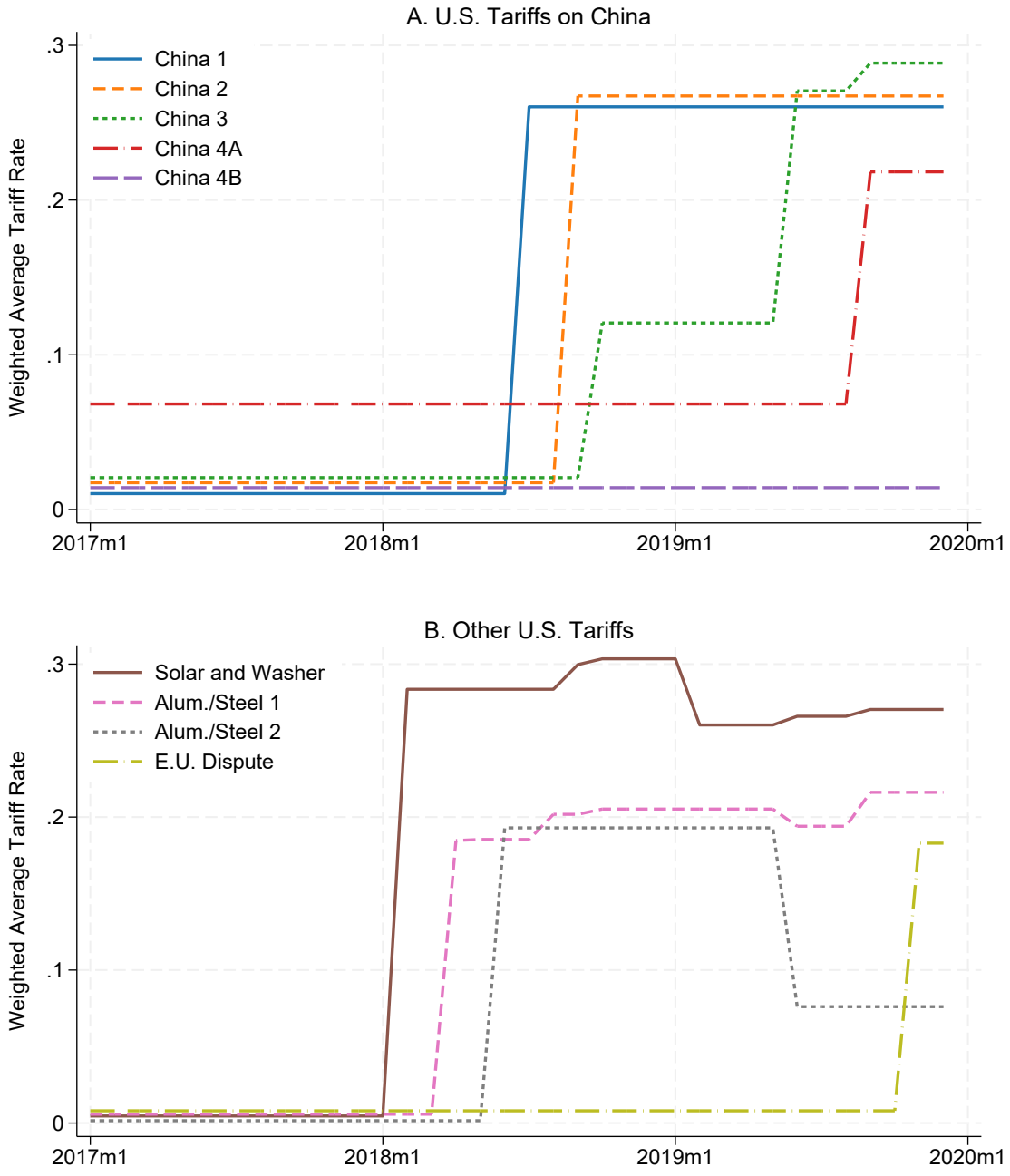
This section documents full details on U.S. import tariffs used in the study, organized by treatment group. The information on these tariffs is obtained directly from official U.S. announcements.

Solar and Washer (Section 201). The U.S. announced tariffs on crystalline silicon photovoltaic cells (Presidential Proclamation 9693) and large residential washers (Presidential Proclamation 9694) imports on January 23, 2018 (Federal Register, Vol. 83, No. 17). They went into effect on February 7, 2018. The tariffs on CSPV cells impacted all non-GSP (Generalized System of Preferences) countries while those on washers impacted all non-GSP countries with Canada being the notable exemption. For solar products, the additional tariff rate was 30%. For washers, the rates were 20% and 50%. Some of these rates were contingent on surpassing an annual quantity threshold. For this study, I assumed the higher rates that would have applied if the thresholds were surpassed. On February 7, 2019, these rates were lowered: 30% to 25%, 20% to 18%, and 50% to 45%. Panel B of Figure A.1 plots the average tariff rate for affected import varieties, weighted by their 2017 import values. Note that the small rate increases in 2018 and 2019 are due to Chinese varieties being hit with additional tariffs from other waves.

Aluminum and Steel 1 (Section 232). The U.S. announced tariffs on aluminum (Presidential Proclamation 9704) and steel (Presidential Proclamation 9705) on March 8, 2018 (Federal Register, Vol. 83, No. 51). They went into effect on March 23, 2018. Initially, only the following countries were exempted: Canada, Mexico, Australia, Argentina, South Korea, Brazil, and the E.U. (Federal Register, Vol.83, No.60). The additional tariff rate on aluminum products was 10% and those on steel products was 25%. The exemption on aluminum imports from South Korea expired on May 1, 2018. However, I include them into this treatment group given their small trade value. The tariff rate on steel products from Turkey was raised to 50% on August 13, 2018 (Federal Register, Vol. 158, No. 83) and then later reduced back to 25% on May 21, 2019 (Federal Register, Vol. 84, No. 98). Panel B of Figure A.1 plots the average tariff rate for affected import varieties, weighted by their 2017 import values.

Aluminum and Steel 2 (Section 232). Following the initial exemptions, only Argentina and Australia were able to obtain permanent exemptions for aluminum products while only Argentina, Australia, Brazil, and South Korea obtained permanent exemptions for steel products (Federal Register, Vol. 83, No. 108). For the other initially exempted import varieties (Canada, Mexico, Brazil (aluminum), the E.U.), the same tariff rates applied

Figure A.1: Monthly U.S. Import Tariff Rates by Treatment Group



Notes: For each treatment group, the product-level tariff rates are averaged using their 2017 import value as weights.

to other countries were applied to them as of June 1, 2018. For Canada and Mexico, these tariffs were lifted on May 20, 2019 (Federal Register, Vol. 84, No. 100). Panel B of Figure A.1 plots the average tariff rate for affected import varieties, weighted by their 2017 import values.

China 1 (Section 301). The initial product list for the first wave of U.S. tariffs on China was released on April 3, 2018 (also see Federal Register, Vol. 83, No. 67). The final list was released on June 20, 2018 and implemented on July 6, 2018 (Federal Register, Vol. 83, No. 119) with an additional tariff rate of 25%.

China 2 (Section 301). The initial product list for the second wave of U.S. tariffs on China was released on June 15, 2018 (also see Federal Register, Vol. 83, No. 119). The final list was released on August 16, 2018 and implemented on August 23, 2018 (Federal Register, Vol. 83, No. 159) with an additional tariff rate of 25%.

China 3 (Section 301). The initial product list for the third wave of U.S. tariffs on China was released on July 10, 2018 (also see Federal Register, Vol. 83, No. 137). The final list was released on September 21, 2018 and implemented on September 24, 2018 (Federal Register, Vol. 83, No. 184) with an additional tariff rate of 10%. The tariff rate was initially scheduled to increase to 25% on January 1, 2019 but was postponed until May 10, 2019 (Federal Register, Vol. 84, No. 90).

China 4A (Section 301). The initial product list for the fourth wave of U.S. tariffs on China was released on May 17, 2019 (Federal Register, Vol. 84, No. 96). The final list was released on August 20, 2019 and implemented on September 1, 2019 (Federal Register, Vol. 84, No. 161) with an additional tariff rate of 15%, which was increased from the initially announced 10% (Federal Register, Vol. 84, No. 169).

China 4B (Section 301). The initial product list for the fifth wave of U.S. tariffs on China was released on May 17, 2019 (Federal Register, Vol. 84, No. 96). These tariffs were scheduled to go into effect on December 15, 2019 with an additional rate of 15% (Federal Register, Vol. 84, No. 169). However, they were suspended following the Phase One deal between the U.S. and China (Federal Register, Vol. 84, No. 243).

Panel A of Figure A.1 plots the average tariff rate for affected import varieties for each of these tariff waves against China, weighted by their 2017 import values.

E.U. Dispute (Section 301). The initial product list for the U.S. tariffs on the E.U. following its dispute over large civil aircraft was released on April 8, 2019 (Federal Register, Vol. 84, No. 71). The final list was released on October 18, 2019 and implemented on the same day with additional rates of 10% and 25% (Federal Register, Vol. 84, No. 202). Panel B of Figure A.1 plots the average tariff rate for affected import varieties, weighted by their 2017 import values.

A.2 Foreign Retaliatory Tariffs

This section documents full details on foreign retaliatory tariffs used in the study, organized by treatment group. The information on these tariffs is obtained directly from respective retaliating countries' official announcements.

Several countries retaliated against the U.S. Section 232 tariffs on aluminum and steel products. Panel B of Figure A.2 plots the average tariff rate for affected export varieties, weighted by their 2017 export values. Here are additional details on these tariffs.

Aluminum and Steel: China. China was first to retaliate. China announced the retaliation on March 23, 2018 and implemented it on April 2, 2018 with additional tariff rates of 15% and 25%.

Aluminum and Steel: Mexico. Mexico announced its retaliation and published a product list on June 5, 2018, which went into effect immediately with additional tariff rates ranging 5-25%. The retaliatory tariffs on a subset of the products appearing on this list were increased on July 5, 2018. These tariffs were lifted on May 20, 2019 following a trade deal with the U.S.

Aluminum and Steel: Europe. The E.U., Turkey, and Russia all retaliated. The E.U. initially released a list of retaliatory tariffs on May 16, 2018, which were revised and implemented on June 22, 2018 with additional tariff rates of 10% and 25%. The proposed list included future rate increases in 2021, which is beyond the scope of this study. Turkey notified the WTO of its intention to retaliate with a proposed list on May 22, 2018. The tariffs were first implemented on June 21, 2018 with additional rates ranging 4-70%. Following the U.S. doubling its tariffs on Turkish steel products, Turkey increased its tariff rates on August 15, 2018, which were later lowered back on May 22, 2019. Russia first released a product list for retaliation on July 6, 2018. These tariffs were implemented on August 5, 2018 with rates ranging 25-40%.[8]

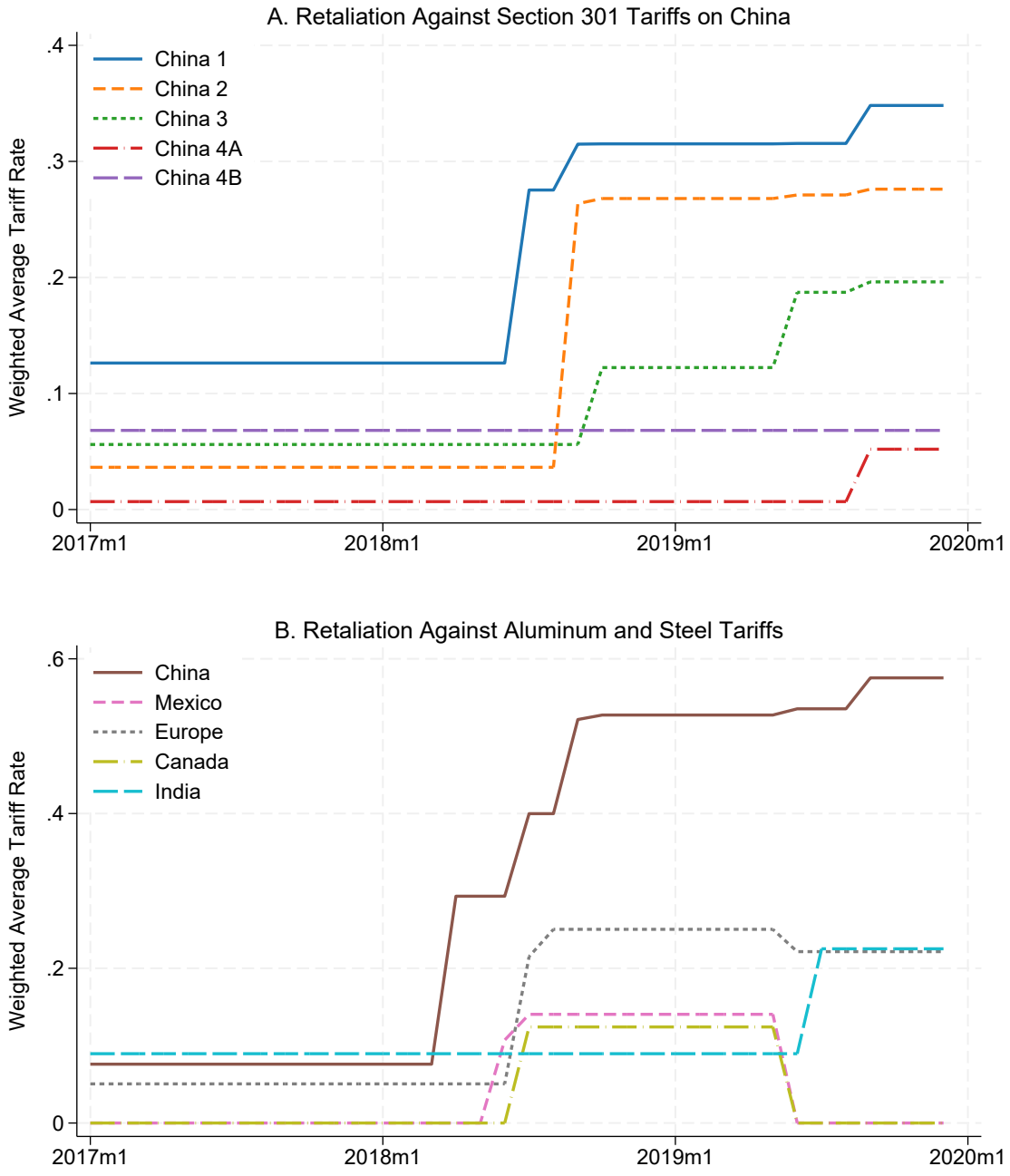
Aluminum and Steel: Canada. Canada announced its retaliation and released a product list on May 31, 2018. These tariffs were revised and implemented on July 1, 2018 with additional tariff rates of 10% and 25%. These tariffs were eventually lifted on May 19, 2019 following a trade deal with the U.S.

Aluminum and Steel: India. India notified the WTO of its intent to retaliate and released a product list on May 18, 2018. Following a series of postponements, the retaliatory tariffs were revised and implemented on June 16, 2019 with additional rates ranging 5-20%.

China announced retaliations against each of the five waves of U.S. Section 301 tariffs it faced. Panel A of Figure A.2 plots the average tariff rate for affected export varieties,

[8]The link to the official announcement no longer works. I have saved a copy of the announcement in the replication package.

Figure A.2: Monthly Foreign Tariff Rates on U.S. Exports by Treatment Group



Notes: For each treatment group, the product-level tariff rates are averaged using their 2017 export value as weights.

weighted by their 2017 export values. Further details on these retaliations are as follows.

China 1. In response to the first wave of U.S. tariffs, China initially released a list of retaliatory tariffs on April 4, 2018, which were revised and implemented on July 6, 2018 with an additional tariff rate of 25%.

China 2. The product list for the retaliation against the second wave was released on June 16, 2018. It was later revised and implemented on August 23, 2018 with an additional tariff rate of 25%.

China 3. For the retaliation against the third wave, the initial product list was released on August 3, 2018. The list was later implemented on September 24, 2018 with additional tariff rates of 5% and 10%, which were revised downward. The originally proposed rates were scheduled to go into effect at the end of the year but were further delayed. Then, following the increase in tariff rates by the U.S., China increased the rates on these products on June 1, 2019 to 5%, 10%, 20%, or 25%, which were the originally proposed rates.

China 4A. The initial product list was released on August 23, 2019 and implemented on September 1, 2019 with additional tariff rates of 5% and 10%.

China 4B. The initial product list was released on August 23, 2019 together with the list for the retaliatory tariffs in China 4A. As with the U.S. tariffs, these retaliatory tariffs were scheduled to go into effect on December 15, 2019 with additional tariff rates of 5% and 10%. However, they were suspended following the Phase One deal between the two countries and never implemented during the sample period.

B Empirical Appendix

B.1 Estimation of Average Pass-Through Rates

This section describes the estimation of average pass-through rates by aggregating the observation-level pass-through rate estimates in equation (19). In particular, I lay out the steps I use to adjust the weights so that the `did_imputation` Stata package of Borusyak et al. (2024) can be used for the estimation and the calculation of standard errors.

$$\begin{aligned}
\widehat{APTR}_t^{pm*} &= 1 + \sum_{g \in G^M} \sum_{ij \in \Omega_g^M} \left(\frac{m_{ij0}}{\sum_{g' \in G^M} \sum_{i'j' \in \Omega_{g'}^M} m_{i'j'0}} \right) \frac{\hat{\theta}_{ijt}^{pm*}}{\Delta \ln(1+\tau_{ijt})} \\
&= 1 + \sum_{g \in G^M} \sum_{ij \in \Omega_g^M} \left(\frac{m_{ij0}/\Delta \ln(1+\tau_{ijt})}{\sum_{g' \in G^M} \sum_{i'j' \in \Omega_{g'}^M} m_{i'j'0}} \right) \hat{\theta}_{ijt}^{pm*} \\
&= 1 + \frac{1}{\widetilde{\tau}_t} \sum_{g \in G^M} \sum_{ij \in \Omega_g^M} \left(\frac{m_{ij0}/\Delta \ln(1+\tau_{ijt})}{\sum_{g' \in G^M} \sum_{i'j' \in \Omega_{g'}^M} m_{i'j'0}/\Delta \ln(1+\tau_{i'j't})} \right) \theta_{ijt}^{pm*},
\end{aligned}$$

where $\widetilde{\tau}_t$ is the weighted harmonic mean of log changes in tariffs defined as

$$\widetilde{\tau}_t = \left[\sum_{g \in G^M} \sum_{ij \in \Omega_g^M} \left(\frac{m_{ij0}}{\sum_{g' \in G^M} \sum_{i'j' \in \Omega_{g'}^M} m_{i'j'0}} \right) \Delta \ln(1+\tau_{ijt})^{-1} \right]^{-1}.$$

Thus, the aggregate APTR can be estimated by using import values divided by tariff changes as weights. Then, the resulting estimate can be divided by the weighted harmonic mean of tariff changes and add one to arrive at the final estimate.

B.2 Aggregation of Estimates by Event Time

The imputed observation-level effects in equation (15) can be reorganized in time relative to their respective initial tariff implementation dates (i.e., event time):

$$\widehat{ATT}_e^{pm*} = \sum_{g \in G^M} \sum_{ij \in \Omega_g^M} \left(\frac{m_{ij0}}{M_{g0}} \right) \hat{\theta}_{ij,T_{ij}^*+e}^{pm*}, \tag{21}$$

where T_{ij}^* denotes the date in which tariffs were first implemented for variety ij. This is useful when evaluating the treatment effect dynamics relative to tariff implementation dates while continuing to rely on February 2018 as the relevant treatment date for identification.