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

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Abstract

This paper takes advantage of recent econometric advancements in the difference-in-differences literature to shed light on why conventional event studies obtain biased estimates for the 2018–2019 trade war's impact on U.S. prices. Accounting for the staggered and heterogeneous impacts on U.S. prices, I estimate that the trade war led both foreign and U.S. exporters to lower their pre-tariff prices in 2019 by 8% and 13%, respectively, after having no initial impacts in 2018. Because the trade war affected a much larger value of U.S. imports than exports, the U.S. experienced a positive terms-of-trade gain equivalent to $9.9 billion in 2019. I also find that the heterogeneity in price responses to the tariffs implies a relatively muted impact on the final import prices of consumer goods. Finally, I show that decreases in pre-tariff prices were limited to products where the importer held a large share of the exporter's market.

1 Introduction

The 2018–2019 trade war between the U.S. and its major trade partners has been the subject of numerous studies. A robust yet surprising finding across a number of research (Amiti et al., 2019, 2020; Fajgelbaum et al., 2020; Cavallo et al., 2021) has been the nearly complete pass-through of U.S. tariffs. In comparison, the evidence on the pass-through of retaliatory tariffs is more mixed (Cavallo et al., 2021; Ma et al., 2021; Jiao et al., 2024). Given the large sizes of the U.S. and China's economy in particular, economists expected the reduced import demand from higher tariffs to induce exporters to lower their prices as conventional trade theories would suggest. However, Amiti et al. (2020) and Cavallo et al. (2021) found nearly complete pass-through of U.S. import tariffs well into 2019, with the former study

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finding incomplete pass-through only on steel inputs.[1] In this paper, I present new event study estimates that indicate incomplete pass-through of both U.S. and retaliatory tariffs by taking advantage of recent econometric advancements in the difference-in-differences (DiD) literature.

The tariffs exchanged during the trade war in 2018–2019 were implemented in a staggered manner, one round of U.S. tariffs typically followed by a round of retaliatory tariffs. However, existing event study estimates do not account for the staggered timing of the tariffs, which the recent DiD literature (de Chaisemartin and D'Haultfœuille, 2020; Goodman-Bacon, 2021; Callaway and Sant'Anna, 2021; Sun and Abraham, 2021; Borusyak et al., 2024) has shown to be problematic. These papers demonstrate, in different ways, how conventional event study specifications that do not account for the staggered timing of treatment can produce biased estimates even when identification assumptions hold perfectly. For example, Sun and Abraham (2021) show that conventional event studies implicitly compare the outcomes of treated units with those of other already-treated units.

A comparison of conventional and staggered event study estimates for the effects of U.S. tariffs on pre-tariff import prices and values confirms these issues. As in previous studies, a conventional event study estimates that pre-tariff import prices did not fall while import values fell by 50% about a year and a half after tariffs were announced. Adapting the estimation procedure proposed in Sun and Abraham (2021) to the current setting, I instead find that pre-tariff import prices fell by 13% and import values by 55% about a year and a half after tariffs were announced. Interestingly, the discrepancy between the two specifications exists only for pre-tariff import prices and not import values. This discrepancy stems from the fact that treatment effects across tariff waves show homogeneous dynamics for import values but not prices. Sun and Abraham (2021) show that treatment effect homogeneity is a special case in which the conventional event study produces unbiased estimates.

Using the staggered event study estimates, I re-evaluate the trade war's impact on U.S. import and export prices. The estimates indicate that pre-tariff prices for both tariff-hit imports and exports fell primarily in 2019. In 2018, their pre-tariff prices showed little change and so the tariff pass-through was nearly complete for both U.S. and retaliatory tariffs, which is consistent with the conclusions in earlier studies (Amiti et al., 2019; Fajgelbaum et al., 2020) that focused on 2018. However, in 2019, pre-tariff prices of targeted imports fell by 8% and those of exports fell more strongly by 13%. As a result, tariff-inclusive prices of U.S. imports rose only by 7% while those of exports did not increase at all. The corresponding tariff pass-through rates were 44% and 0%, respectively.

[1] In an unpublished follow-up manuscript, Fajgelbaum et al. (2020) also show continued complete pass-through of both U.S. and retaliatory tariffs in 2019.

Combining these new estimates with a standard framework (Dixit and Norman, 1980), I assess the 2018–2019 trade war's impact on U.S. welfare. For 2018, I estimate a negligible welfare impact due to a lack of pre-tariff price responses for both imports and exports. However, for 2019, I estimate a welfare gain of 0.05% or $9.9 billion. The foreign exporters' willingness to lower their prices led to a terms-of-trade gain for the U.S. Although U.S. exporters lowered their prices by greater margins than foreign exporters, the net terms-of-trade effect was positive due to the fact that the trade war affected a much larger value of U.S. imports than exports.

These aggregate price effects, however, mask substantial heterogeneity across tariff waves that have important implications on U.S. prices. I find that increases in tariff-inclusive import prices were higher for the earlier waves of tariffs, which primarily landed on non-consumer goods (e.g., intermediate inputs and capital goods). Consequently, I estimate that tariff-inclusive prices of consumer goods rose by only 4%, which is half the increase estimated for non-consumer goods. This result offers a complementary explanation as to why Cavallo et al. (2021) found little evidence of higher U.S. retail prices resulting from the trade war. Its impact on the import prices of consumer-facing goods was more muted. In comparison, the retaliatory tariffs had equal impacts across consumer and non-consumer goods.

I also explore potential sources underlying the heterogeneous price responses. Trade theory suggests that tariffs levied by large importing countries can induce foreign exporters to lower their prices. I test this idea empirically by first measuring each importing country's market share of its trade partner's export market for each product. This bilateral measure is close to one, for example, if U.S. imports take up a large share of China's total exports of the product. In accordance with the trade theory's prediction, I find that both U.S. and foreign exporters lowered their prices primarily for products where the importing country held high market shares.

The results in this paper highlight the shortcomings of conventional event study designs within the context of the 2018–2019 trade war and suggest that the literature's assessment of the trade war's impact may be incomplete. The staggered event study estimates indicate that the trade war's impact on U.S. prices is much more in line with the predictions of standard trade theory. As the U.S. is going through another trade war and numerous studies about its impacts are sure to follow, the findings of this paper should caution future studies of the trade war from relying on conventional event study designs.

2 Theoretical and Empirical Framework

This section lays down the theoretical framework for assessing the welfare impacts of both U.S. and retaliatory tariff changes. The framework will clarify which objects need to be estimated, and how they should be aggregated to calculate the welfare impact.

2.1 Theoretical Framework

To preface the discussion on identification, I follow the notations in Baker et al. (2025) and allow potential outcomes to vary by treatment groups. Let Ω^M and Ω^X denote the set of imported and exported varieties, respectively. An imported (exported) variety refers to a product i sourced from (shipped to) country j. Imported varieties are partitioned into different treatment groups $g \in G^M$, where G^M is the set of all treatment groups. I use G_{ij} to denote the treatment group of variety ij and assign $G_{ij} = \infty$ for never-treated varieties. Thus, the set of varieties in group g is given by $\Omega^M_g = \{ij \in \Omega^M | G_{ij} = g\}$ and $\Omega^M = \cup_{g \in G^M} \Omega^M_g$. These concepts are analogously defined for exported varieties using the superscript X instead of M. Then, the observed import value of variety ij at time t, for example, can be written as

$$m_{ijt} = \sum_{g \in G^M} m_{ijt}(g) 1\{G_{ij} = g\}, \tag{1}$$

where $m_{ijt}(g)$ denotes the potential outcome from receiving the treatment that varieties in group g received and $1\{G_{ij} = g\}$ is a indicator variable that equals one if and only if $G_{ij} = g$.

I assess the welfare impact of observed tariff changes at time t by its equivalent variation. This is the hypothetical change in domestic income at the counterfactual equilibrium (where there are no tariff changes) that would have an equivalent welfare effect as the tariff changes. Dixit and Norman (1980) show that the first-order approximation of the equivalent variation at time t can be calculated as

$$EV_t = \underbrace{\sum_{ij \in \Omega^M} r_{ijt}(\infty) \Delta \ln q^m_{ijt}}_{\text{tariff revenue effect}} \underbrace{- \sum_{ij \in \Omega^M} m_{ijt}(\infty) \Delta \ln p^{m*}_{ijt} + \sum_{ij \in \Omega^X} x_{ijt}(\infty) \Delta \ln p^{x*}_{ijt}}_{\text{terms of trade effect}}, \tag{2}$$

where $r_{ijt} = \tau^m_{ijt} m_{ijt}$ is the tariff revenue on imports; 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; and τ^m_{ijt} are ad-valorem import tariffs such that $p^m_{ijt} = (1 + \tau^m_{ijt}) p^{m*}_{ijt}$ are the prices of imports that domestic consumers face. The difference operator Δ is defined with respect to the counterfactual equilibrium such that $\Delta \ln p^{m*}_{ijt} = \ln p^{m*}_{ijt} - \ln p^{m*}_{ijt}(\infty)$, for example.

The first two terms in equation (2) capture the net effect of import tariff changes on domestic consumers, producers, and government revenue. The equation further 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). Similarly, the third term captures the effect of retaliatory tariff changes on domestic exporters, which will be more negative if foreign tariffs induce domestic exporters to lower pre-tariff prices (i.e., $\Delta \ln p_{ijt}^{x*} < 0$). 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.

Using equation (1) and annual trade values from 2017 (denoted as r_{ij0}, m_{ij0}, and x_{ij0}, respectively) in lieu of the unobserved values $r_{ijt}(\infty)$, $m_{ijt}(\infty)$, and $x_{ijt}(\infty)$ in equation (2), I can obtain the following equivalent variation:

$$EV_t^0 = \underbrace{\sum_{g \in G^M} R_{g0} \cdot ATT_{gt}^{qm}}_{\equiv QM_t^0} + \underbrace{\left(- \sum_{g \in G^M} M_{g0} \cdot ATT_{gt}^{pm*} \right)}_{\equiv PM_t^0} + \underbrace{\sum_{g \in G^X} X_{g0} \cdot ATT_{gt}^{px*}}_{\equiv PX_t^0}, \tag{3}$$

where $R_{g0} = \sum_{ij \in \Omega_g^M} r_{ij0}$, $M_{g0} = \sum_{ij \in \Omega_g^M} m_{ij0}$, $X_{g0} = \sum_{ij \in \Omega_g^X} x_{ij0}$, and the (weighted) average treatment effect on treated units (ATT) terms are given by

$$ATT_{gt}^{qm} = \mathbb{E}_r \left[\Delta \ln q_{ijt}^m(g) \middle| G_{ij} = g \right] = \frac{\mathbb{E}\left[r_{ij0} \Delta \ln q_{ijt}^m(g) \middle| G_{ij} = g \right]}{\mathbb{E}\left[r_{ij0} | G_{ij} = g \right]} \tag{4}$$

$$ATT_{gt}^{pm*} = \mathbb{E}_m \left[\Delta \ln p_{ijt}^{m*}(g) \middle| G_{ij} = g \right] = \frac{\mathbb{E}\left[m_{ij0} \Delta \ln p_{ijt}^{m*}(g) \middle| G_{ij} = g \right]}{\mathbb{E}\left[m_{ij0} | G_{ij} = g \right]} \tag{5}$$

$$ATT_{gt}^{px*} = \mathbb{E}_x \left[\Delta \ln p_{ijt}^{x*}(g) \middle| G_{ij} = g \right] = \frac{\mathbb{E}\left[x_{ij0} \Delta \ln p_{ijt}^{x*}(g) \middle| G_{ij} = g \right]}{\mathbb{E}\left[x_{ij0} | G_{ij} = g \right]} \tag{6}$$

These ATT terms cannot be directly measured since the outcome values under no-treatment are unobserved for treated varieties. It will require additional assumptions to estimate them.

2.2 Empirical Specification

Previous studies (Fajgelbaum et al., 2020; Amiti et al., 2020; Ma et al., 2021; Jiao et al., 2024) have used variations of the following pooled event study specification to estimate the trade war's impact on some outcome variable y_{ijt}:

$$\ln y_{ijt} = \alpha_{ij} + \delta_t + \sum_{s \neq -1} \beta_s \cdot 1\{t = T_{ij} + s\} + \epsilon_{ijt}, \tag{7}$$

where α_{ij} is the variety fixed effect, δ_t is the time fixed effect, and ϵ_{ijt} is the error term. $1\{t = T_{ij} + s\}$ is an indicator variable that equals one if t is s months from T_{ij}, the time in which variety ij was first treated during 2018–2019, and zero otherwise. For example, treatment could refer to the announcement of tariffs. Typically, the month prior to the treatment date ($t = T_{ij} - 1$) is omitted to avoid multi-collinearity and is used as the reference pre-treatment period. This allows for a difference-in-differences (DiD) interpretation of the parameter β_s. It can be interpreted as measuring the difference in outcomes between the treated and untreated varieties, s months from the treatment relative to the month prior to the treatment (i.e., the omitted period).

The DiD interpretation makes it convenient for researchers to discuss identification assumptions using the familiar languages of no-anticipation and parallel trends. The "No Anticipation" assumption can be stated as follows.

Assumption 1 (No Anticipation)**.** *For every eventually treated group $g \in \Omega_g^M$ (or Ω_g^X), outcome variable y_{ijt} satisfies $y_{ijt}(g) = y_{ijt}(\infty)$ for all varieties in group g and pre-treatment periods $t < T_{ij}$.*

This assumption says that treatment effects should not materialize prior to the treatment date so that outcomes in pre-treatment periods are truly untreated. The second assumption is the "Parallel Trends" assumption and can be stated as follows.

Assumption 2 (Parallel Trends)**.** *For every treatment group $g, g' \in \Omega_g^M$ (or Ω_g^X) and periods t, the outcome variable y_{ijt} and corresponding weights $\omega = \{\omega_{ij0}\}$ satisfy*

$$\mathbb{E}_\omega \left[y_{ijt}(\infty) - \ln y_{ijt-1}(\infty) | G_{ij} = g \right] = \mathbb{E}_\omega \left[y_{ijt}(\infty) - y_{ijt-1}(\infty) | G_{ij} = g' \right] \tag{8}$$

This parallel trends assumption says that the observed difference in *untreated* outcomes across any treatment groups in any pre-treatment period can be used to measure their unobserved difference in *untreated* outcomes in a post-treatment period. Thus, when these assumptions are satisfied, one may think that β_s in equation (7) summarizes the relevant ATTs of interest in equation (3).

Recently, several papers (de Chaisemartin and D'Haultfœuille, 2020; Goodman-Bacon, 2021; Callaway and Sant'Anna, 2021; Sun and Abraham, 2021; Borusyak et al., 2024) have demonstrated potential issues with the specification in equation (7) when treatment occurs in a staggered manner across different groups. Even when the two identification assumptions hold, the specification in equation (7) can be problematic, because it makes comparisons between outcomes of treated units with those of already-treated units. Sun and Abraham (2021) show that the pooled event study is valid only when treatment effect dynamics are

homogeneous across treatment groups. These authors have proposed different DiD estimators for the ATTs in a setting where treatments are staggered and their effects are heterogeneous.

For this study, I focus on the estimation procedure of Sun and Abraham (2021), because it is particularly well-suited for making comparisons with the pooled event study. Sun and Abraham (2021) proposes estimating the following staggered event study specification:

$$\ln y_{ijt} = \alpha_{ij} + \delta_t + \sum_{g \neq \infty} \sum_{s \neq -1} \beta_{gs} \cdot 1\{G_{ij} = g\} \cdot 1\{T_{ij} + s = t\} + \epsilon_{ijt}, \tag{9}$$

where the never-treated units ($g = \infty$) are the omitted reference treatment group. This specification differs from the pooled one in equation (7) in that it interacts all relative time and treatment group indicators, only omitting the periods immediately before the treatment dates.

The parameter β_{gs} has a 2x2 DiD interpretation that clarifies the assumptions required for its causal interpretation. It measures the difference in outcomes between units in group g and the never-treated group, s months from the treatment of units in group g relative to the month prior to the treatment. As Baker et al. (2025) clarifies, the necessary parallel trends assumption in this case is given by

Assumption 3 (Parallel Trends, SA)**.** *For every eventually treated group $g \neq \infty$ and post-treatment periods $t \geq T_g$, the outcome variable y_{ijt} and corresponding weights $\omega = \{\omega_{ij0}\}$ satisfy*

$$\mathbb{E}_\omega \left[y_{ijt}(\infty) - \ln y_{ijt-1}(\infty) | G_{ij} = g \right] = \mathbb{E}_\omega \left[y_{ijt}(\infty) - y_{ijt-1}(\infty) | G_{ij} = \infty \right] \tag{10}$$

Thus, trends in the untreated outcomes of units in group g and the never-treated group must be parallel to one another from at least one period prior to the treatment date for group g (T_g). Under Assumptions 1 and 3, β_{gs} identifies the relevant ATTs in equation (3). In practice, researchers typically include additional fixed effects into the specification, but the arguments in Sun and Abraham (2021) still hold when working with variables that have been partialled out on some set of fixed effects.

3 Data and Treatment Groups

The monthly (t) U.S. import and export flows covering 2017–2019 are from the U.S. Census, each reported at their respective 10-digit product codes (i) and by trading country (j). Unless noted otherwise, I use data from the imports for consumption (as opposed to general imports) and domestic exports (as opposed to total exports). The data records both the

value and quantity (when applicable) of imported and exported varieties, which are used to calculate the pre-tariff prices as unit values. For imports, the data also reports estimated collected duties. These trade flows are then matched to their corresponding import tariffs levied by the U.S. and its trade partners during 2017–2019. The matching is performed at the 10-digit product level for imports and at the 6-digit for exports.

The U.S. import tariff changes during 2018–19 are obtained directly from the relevant U.S. Federal Register while the initial import tariff rates in 2017 are calculated by dividing the annual value of collected duties by the customs value of imports. These tariff rates are directly mapped to the relevant data on U.S. import varieties. I obtain the retaliatory tariff changes during 2018–2019 from Fajgelbaum et al. (2024), which are typically recorded using the reporting country's tariff lines (8-digit product codes). Since product codes are internationally consistent only up to the first six digits, I first need to aggregate them by 6-digit product codes before matching them with data on U.S. exports. Thus, I begin with the initial tariff rates in 2017 that were applied on U.S. exports by the retaliating countries. These are the 2017 applied MFN rates obtained from the World Bank's WITS (World Integrated Trade Solution) and are recorded using the reporting countries' tariff lines.[2] Subsequent retaliatory tariff changes during 2018–2019 are added to these rates in 2017 and then averaged by 6-digit product code and retaliating country. The resulting panel of retaliatory tariff rates is then mapped to the data on U.S. exports.[3]

Table 1 lists the different waves of tariffs based on their initial implementation dates.[4] Relative to Amiti et al. (2020) whose data covered until October 2019, the data in this study additionally covers the tariff wave concerning U.S. disputes with the E.U. over aircraft subsidies and two more months of import data but remain largely similar. Table 1 shows the 2018–2019 trade war's disproportionate impact on U.S. imports from and exports to China, especially waves 3 and 4A for imports and waves 1 and 3 for exports. Thus, the treatment effects identified on these waves will matter most significantly for the welfare assessment. On the other hand, the tariff revenues collected by the U.S. were quite small due to the fact that

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

[3]Note that the retaliatory tariffs will be measured less accurately than the import tariffs for two reasons. First, the retaliatory tariffs are mapped to U.S. trade data at the 6-digit product code. Thus, all varieties exported to a retaliating country that share the same first six digits with varieties that are being hit with retaliatory tariffs will be considered to be hit as well. Second, the average retaliatory tariffs at 6-digit product codes are not weighted by trade value, which would require additional data on retaliating countries' imports from the U.S. for each 8-digit product code.

[4]To reduce computational burden and given their small trade values, I group a few tariff waves into larger treatment groups. For imports, the tariffs on aluminum products from South Korea went into effect on May 1, 2018 after a brief exemption but are grouped into "Aluminum and Steel 1." For exports, retaliatory tariffs by Turkey (June 21, 2018), the E.U. (June 22, 2018), and Russia (August 6, 2018) are grouped into "Alum. and Steel: Europe."

Table 1: Summary Statistics by Treatment Group

Panel A: U.S. Import Tariffs

Treatment Group	First Implemented	First Announced	Import Value	Tariff Rate (%) 2017	2018	2019	Consumer Good
Solar and Washer	Feb 07, 2018	Jan 23, 2018	9,011	0.5	30.3	27.0	23%
Alum. and Steel 1	Mar 23, 2018	Mar 08, 2018	18,199	0.6	20.5	21.6	0%
Alum. and Steel 2	Jun 01, 2018	Mar 08, 2018	22,912	0.2	19.3	7.6	0%
China 1	Jul 06, 2018	Apr 03, 2018	31,864	1.0	26.0	26.0	1%
China 2	Aug 23, 2018	Jun 15, 2018	13,259	1.7	26.7	26.7	7%
China 3	Sep 24, 2018	Jul 10, 2018	186,694	2.1	12.1	28.9	32%
China 4A	Sep 01, 2019	May 17, 2019	98,884	6.8	6.8	21.8	72%
China 4B	∞	May 17, 2019	151,221	1.4	1.4	1.4	69%
E.U. Dispute	Oct 18, 2019	Apr 08, 2019	10,072	0.8	0.8	18.3	47%
Never Treated	∞	∞	1,785,037	1.1	1.1	1.1	28%
Total			2,327,153	1.4	3.2	5.1	32%

Panel B: Retaliatory Tariffs

Treatment Group	First Implemented	First Announced	Export Value	Tariff Rate (%) 2017	2018	2019	Consumer Good
Alum. and Steel: China	Apr 02, 2018	Mar 23, 2018	2,443	7.6	52.7	57.5	43%
Alum. and Steel: Mexico	Jun 05, 2018	Jun 05, 2018	6,339	0.0	14.7	0.0	43%
Alum. and Steel: Europe	Jun 21, 2018	May 16, 2018	8,251	4.4	26.9	24.4	48%
Alum. and Steel: Canada	Jul 01, 2018	May 31, 2018	16,594	0.0	12.4	0.0	48%
Alum. and Steel: India	Jun 16, 2019	Jun 20, 2018	1,260	9.0	9.0	24.6	67%
China 1	Jul 06, 2018	Apr 04, 2018	27,806	12.6	31.5	34.8	57%
China 2	Aug 23, 2018	Jun 16, 2018	11,661	3.6	26.8	27.6	2%
China 3	Sep 24, 2018	Aug 03, 2018	43,669	5.6	12.2	19.6	10%
China 4A	Sep 01, 2019	Aug 23, 2019	5,098	0.7	0.7	5.2	14%
China 4B	∞	Aug 23, 2019	1,181	6.8	6.8	6.8	10%
Never Treated	∞	∞	1,184,093				19%
Total			1,308,395				20%

Notes: Trade values are for 2017 and reported in millions of U.S. dollars based on "Imports: For Consumption (Customs Value)" and "Exports: Domestic" from the U.S. Census Bureau. The symbol ∞ indicates that the implementation or announcement of tariffs on the varieties in the relevant group never occurred during 2018–2019. The end-of-year tariff rates are calculated by averaging product-country level (10-digit products for imports and 6-digit for exports) tariff rates weighted by import or export values in 2017. The retaliatory 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 Panel and Washer" products are conditional on being over a quantity threshold and will be lower if under the threshold. The exemption on aluminum products from South Korea expired on Apr 30, 2018 but affected varieties are grouped into "Aluminum and Steel 1". Export varieties subject to retaliatory tariffs from Turkey (Jun 21, 2018), the E.U. (Jun 22, 2018), and Russia (Aug 6, 2018) are group together as "Aluminum and Steel: Europe". The last column reports the 2017 value share of consumer goods (U.S. Census end-use codes 0 or 4) within the treatment group.

the trade-weighted average import tariff rate was only 1.4% in 2017. Thus, we can already anticipate a small welfare impact of tariff revenue effects (QM_t^0) in equation (2).

For each tariff wave or treatment group, I also list the corresponding initial announcement date. These are when the initial product lists of the eventually implemented tariffs were released by the U.S. or a retaliating country. The initial proposals were typically released a few months earlier to collect public comments and requests for exemption. Given the time gap between the announcement and implementation of tariffs, the no-anticipation assumption will likely be violated if the tariff implementation date is used as the treatment date. While pinpointing when importers and exporters could have anticipated future tariff waves is challenging, they would have certainly been able to do so by the time the initial proposals were publicly released. Thus, I use these announcement dates as the treatment dates. When converting the tariff implementation and announcement dates to year-months, I follow Fajgelbaum et al. (2020) and assign the same month if it falls on the first 15 days of the month. Otherwise, I assign the next month.

One notable detail concerns the exchange of China 4B tariffs, which were scheduled to go into effect on December 15, 2019. With the agreement over the Phase One trade deal between the two countries on December 13, however, these tariffs were called off. Despite never being implemented, they were announced on May 17, 2019 for the U.S. and August 23, 2019 for China along with detailed product lists. While it will not be possible to examine the effects of tariff implementation for these waves, it would be remiss to consider them as untreated. Therefore, I assign them to their respective treatment groups. With these varieties taken out from the never-treated group, the import and export varieties from and to China that remain in the never-treated group account for 3.4% and 23.5% of all Chinese varieties based on 2017 trade values, respectively.

The final column of Table 1 reports the 2017 value share of consumer goods within each treatment group. The U.S. Census classifies each 10-digit product into one of several categories based on its principal use. For this study, consumer goods refer to the combined categories for "Foods, Feeds, & Beverages" and "Consumer Goods" (end-use codes 0 or 4). To name a few example, "Foods, Feeds, & Beverages" contains bakery products, wheat, and alcoholic beverages while "Consumer Goods" contains apparel, toys, and furniture. The earlier U.S. import tariff waves mostly targeted non-consumer goods with consumer goods showing up in more significant quantities starting with China-3. For retaliatory tariffs, the earlier waves hit all goods more evenly with later waves hitting mostly non-consumer goods. Thus, heterogeneity in the treatment effects across groups will have different implications for the prices of consumer and non-consumer goods.

Finally, to guard against extreme measurement errors in prices (unit values), I do not use

the reported quantity information if the official unit of measurement for the 10-digit product code changed within the sample period. In addition, I first compute the ratio between an observation's unit value and the median unit value for the same variety within the sample. Then, I do not use the reported quantity information of observations with ratios that are in the top and bottom 1 percentiles of each sample period. These cutoffs approximately correspond to unit values that are 25 times larger or smaller than the median unit value of the same variety.

4 Main Results

4.1 Comparison of Pooled and Staggered Event Study Estimates

I begin with a comparison of treatment effects on pre-tariff import prices and import values estimated from pooled and staggered event study specifications. This comparison will be extremely useful in empirically highlighting the shortcomings of the pooled event study specification in this setting. Specifically, I estimate

$$\ln y_{ijt} = \alpha_{ij} + \theta_{it} + \eta_{jt} + \sum_{s \neq -1} \beta_s \cdot 1\{t = T_{ij} + s\} + \epsilon_{ijt}, \tag{11}$$

where θ_{it} and η_{jt} are newly added fixed effects that respectively absorb product and country specific time-varying trends. For instance, demand shocks for air conditioners in the summer months would be absorbed by the product-time fixed effects while changes in exchange rates would be absorbed by the country-time fixed effects. T_{ij} is the treatment time for variety ij, which is the initial announcement date listed in Table 1.

I estimate equation (11) using logged pre-tariff import prices and import values as the outcome variables and weighting by annual import values in 2017. Standard errors are clustered by 8-digit product codes and by country. The resulting estimates are plotted in Figure 1 Panel A and B, focusing on estimates within 17 months from the initial announcement dates.[5] The plotted estimates reproduce well-documented patterns. For pre-tariff import prices, the estimated effects stay mostly flat over time. While there is some declining trend past month seven, it is imprecisely estimated and not enough to conclude anything other than that pre-tariff import prices remained unchanged. In comparison, the effects on import values are precisely estimated and sensible. There is an initial surge in import values following the tariff announcements, which is explained by exporters front-loading their shipments to beat the tariff implementation dates. Afterwards, once the tariffs are implemented, there

[5]This is the latest month that covers all import tariff waves from 2018.

is a precipitous decline in import values. Seventeen months after the tariff announcements, the point estimates for pre-tariff import prices and import values are -0.024 and -0.688, respectively. That is, pre-tariff import prices fell by 2.4% (without statistical significance) and import values fell by 49.7%.

I compare these pooled event study estimates with those from staggered event study specifications. Specifically, I estimate:

$$\ln y_{ijt} = \alpha_{ij} + \theta_{it} + \eta_{jt} + \sum_{g \neq \infty} \sum_{s \neq -1} \beta_{gs} \cdot 1\{G_{ij} = g\} \cdot 1\{T_{ij} + s = t\} + \epsilon_{ijt}, \tag{12}$$

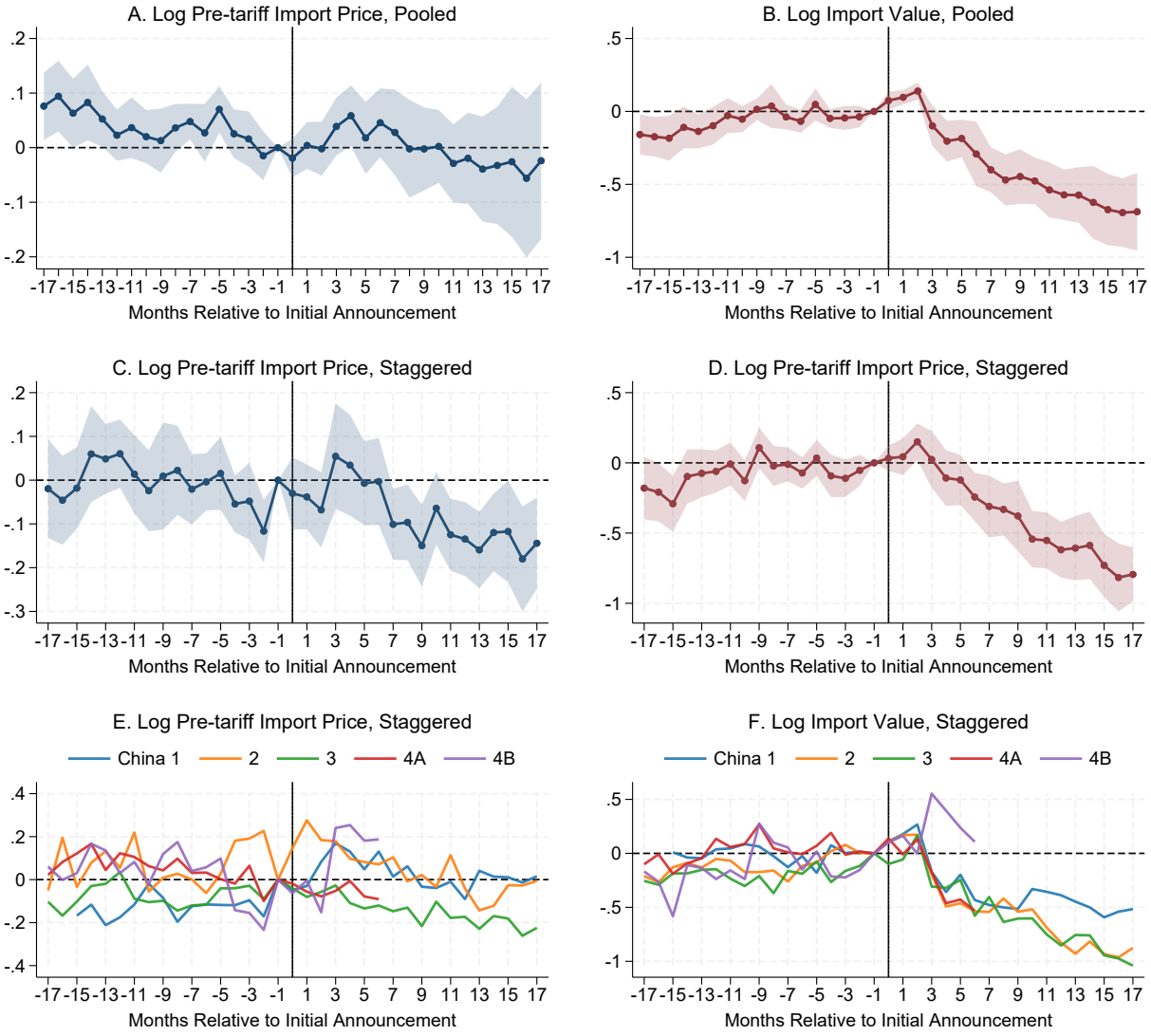
which like equation (11) further introduces product-time and country-time fixed effects. As with the pooled specification, logged pre-tariff import prices and import values are weighted by their annual import values in 2017. Standard errors are clustered by 8-digit product codes and by country. The omission of never-treated group ($g = \infty$) indicates that the varieties in this group are being used as the comparison group for all other treated groups. β_{gs} thus compares outcomes in group g and the never-treated group, s months from the tariff announcement relative to their outcomes during the month prior to the announcement. As discussed in Section 2.2, a causal interpretation of β_{gs} requires no-anticipation assumption to hold and parallel trends to start holding from at least one month prior to the treatment.

The staggered event study specification produces event study estimates for each treatment group and the number of estimates can be rather large. To make a more direct comparison with the pooled event study plots in Panel A and B, I aggregate the staggered event study estimates by event time (i.e., months relative to initial announcement) using their 2017 import values as weights:

$$\hat{\beta}_s = \sum_{g \neq \infty} \frac{M_{g0}}{\sum_{g \neq \infty} M_{g0}} \hat{\beta}_{gs}. \tag{13}$$

Thus, $\hat{\beta}_s$ measures the trade-weighted ATT across all treatment groups, s months from their respective treatment dates. The trade-weighted ATTs are plotted in Figure 1 Panel C and D along with their 95% confidence intervals. The trade-weighted ATTs for import prices in Panel C differ substantially from the estimates in Panel A. The pre-tariff import prices show statistically significant drops from month seven and persist for the next ten months. The point estimate is -0.144 or -13.4% seventeen months from the initial announcement. On the other hand, the trade-weighted ATTs for import values in Panel D display very similar dynamics to the estimates in Panel B. The point estimate is -0.794 or -54.8% seventeen months after the initial announcement, which is similar in magnitude to the pooled event study estimate from earlier.

Figure 1: Comparison of Pooled and Staggered Event Study Estimates



Notes: Panel A and B plot coefficients from estimating the pooled event study specification in equation (11). Panel C and D plot the import-value weighted averages of coefficients from the staggered event study specification in equation (12). Panel E and F plot coefficients from the staggered event study specification for treatment groups associated with U.S. tariffs on China. All figures only show estimates for 17 months from the initial announcement dates. Shaded regions are 95% confidence intervals based on standard errors clustered by product (8-digit) and by country.

A rather deceiving aspect of the pooled event study specification in this setting is that it produces sensible and precise estimates for import values. Given that import prices are simply import values divided by quantities, it appears unlikely that the same specification would successfully identify treatment effects for one variable but not the other. So, how does the pooled event study specification produce biased estimates only for import prices? This discrepancy can be explained by a result in Sun and Abraham (2021). In a general setting, Sun and Abraham (2021) derive that the pooled event study specification can place negative weights on group-specific ATTs and produce biased estimates when treatments are staggered and their effects are heterogeneous. Interestingly, they find that treatment effect homogeneity is a special case in which the pooled event study specification produces unbiased estimates. Could this be the case for treatment effects on import values?

Indeed, I find that the treatment effect dynamics on import prices are quite heterogeneous while those on import values are not. To show this empirically, I plot the estimates of β_{gs} for the five waves of Section 301 tariffs against China in Figure 1 Panel E and F. There is a striking contrast between the ATTs for import prices and values with regards to their heterogeneity across treatment groups. For import prices, there is substantial heterogeneity. The pre-tariff prices of import varieties hit by the first and second waves (China 1 and 2), for example, show large increases following the announcement date before gradually decreasing. The third wave (China 3) shows steady declines following the announcement. The final two waves that were announced in 2019 (China 4A and 4B) interestingly show declining trends leading up to their announcement dates that continue after their tariff announcements, perhaps indicating that they may have been anticipated in 2018.

With import values, the trends are quite homogeneous. They are uniformly flat or slightly upward-trending leading up to the announcements, increase right after, and fall once the tariffs go into effect. Only the canceled China 4B wave departs from these pattern, which justifies separating it out from the never-treated group. Thus, the discrepancy between pooled and staggered event study estimates can be explained by the difference in treatment effect heterogeneity for import prices and values.

4.2 Effects on U.S. Import and Export Prices

In this section, I use the staggered event study estimates to reassess the 2018–2019 trade war's effects on U.S. import and export prices. The estimates in the previous section revealed that monthly trade data, particularly price data, can be noisy. This raises concerns that using a single month as a reference period for each treatment group may lead to less robust estimates. To address this concern, I instead use the six months leading up to tariff announcements as

the reference period. This amounts to further omitting terms that interact treatment group indicators with relevant event time indicators from equation (12), which results in estimating the following specification:

$$\ln y_{ijt} = \alpha_{ij} + \theta_{it} + \eta_{jt} + \sum_{g \neq \infty} \sum_{s \notin [-6,-1]} \mu_{gs} \cdot 1\{G_{ij} = g\} \cdot 1\{T_{ij} + s = t\} + \epsilon_{ijt}, \tag{14}$$

where μ_{gs} compares outcomes in group g and ∞ after s months from tariff announcements relative to their mean outcomes during the six months leading up to the announcements. Hence, relative to β_{gs} in equation (12), a causal interpretation of μ_{gs} additionally requires parallel trends to start holding from at least six months prior to the announcements.

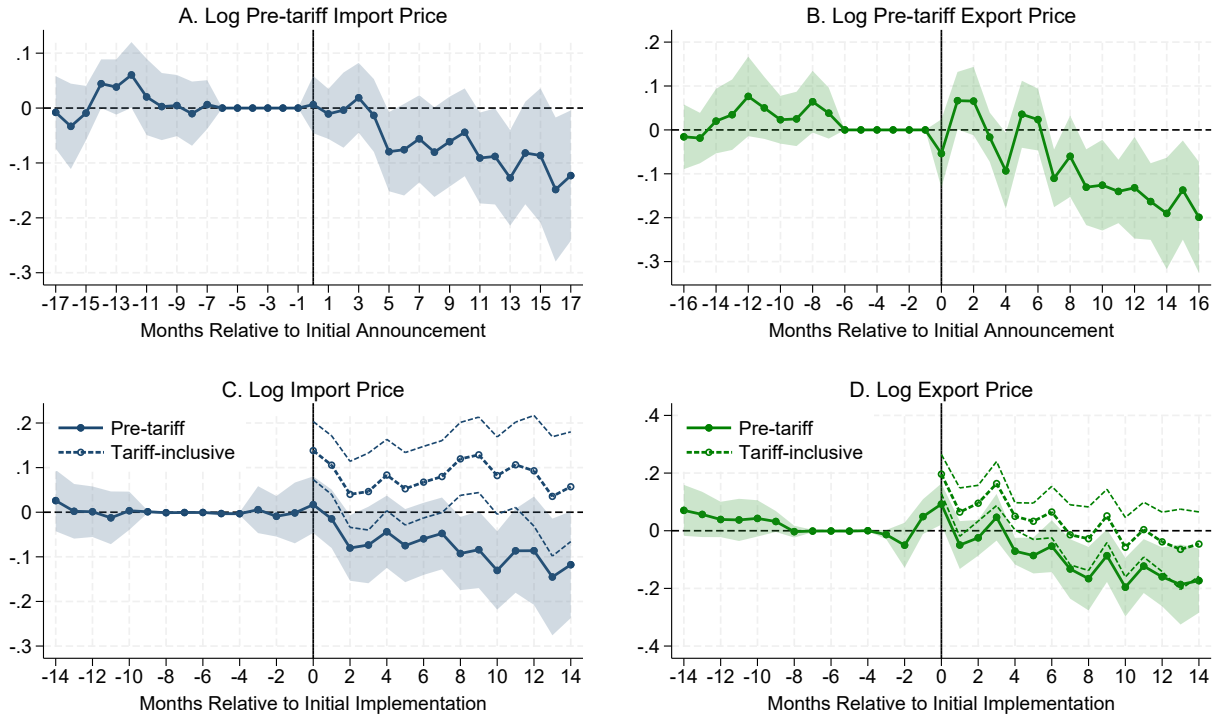
The aggregated ATTs for pre-tariff import and export prices are plotted in Panel A and B of Figure 2. I report the event study plots for individual treatment groups in Section A.1. Import and export varieties in group China 4B are included in the estimation but excluded from the aggregation since their tariffs were called off and never implemented during the sample period. For the other treatment groups, I find that the estimated coefficients in pre-treatment periods are on average not statistically different from zero in all but two cases: Solar and Washer for imported varieties and China 4A for exported varieties. Thus, I exclude these two groups from the aggregation as well but note that these tariff waves represent very small trade values (see Table 1).

Panel A and B of Figure 2 indicate broadly similar patterns for both pre-tariff prices. The ATTs stay close to or fluctuate around zero for a few months after the tariffs are announced but show visible and persistent drops after about half a year. The estimates are -0.123 or -11.6% seventeen months from announcement for import prices and -0.199 or -18.0% sixteen months from announcement for export prices. Like before, these event time horizons were chosen to include all tariff waves from 2018. These results indicate statistically and economically significant lowering of pre-tariff prices for both U.S. imports and exports, with larger effects identified on the latter. The larger effects for U.S. export prices is consistent with the findings in Cavallo et al. (2021).

The estimated coefficient $\hat{\mu}_{gs}$ can be rearranged and aggregated by time relative to the tariff implementation instead.[6] I present these alternatively aggregated trade-weighted ATTs for pre-tariff and tariff-inclusive import prices in Panel C of Figure 2. Analogous plots for export prices are presented in Panel D. These plots give us a better sense on how much of the tariffs were passed onto the importers as opposed to being absorbed by the exporters through lower pre-tariff prices. Panel C shows that tariff-inclusive import prices increased

[6]Note that for some calendar time t, we have $\hat{\mu}_{g,s=t-T_g}$, where T_g is the treatment date of varieties in group g.

Figure 2: Estimated Effects on U.S. Import and Export Prices



Notes: Enclosed regions around point estimates are 95% confidence intervals based on standard errors clustered by product (8-digit for imports and 6-digit for exports) and by country.

significantly and remained steady for most periods despite increases in tariff rates in 2019. The plot reveals that additional tariff rate increases did not result in higher final import prices due to steady declines in pre-tariff import prices. Panel D shows stronger signs that U.S. exporters absorbed most of the retaliatory tariffs by lowering their prices. A year since implementation, the ATT on tariff-inclusive prices are close to zero. However, the results on tariff-inclusive export prices should be read with greater caution since the retaliatory tariffs are only measured at the 6-digit product level and subject to greater measurement error.

A useful concept that summarizes these results is the tariff pass-through rate (ξ), which can be calculated by dividing the estimate for tariff-inclusive prices by the difference between the estimates for tariff-inclusive and pre-tariff prices. For import prices, we have

$$\xi_{gt}^{m} = \frac{\hat{\mu}_{gt}(\ln p_{ijt}^{m})}{\hat{\mu}_{gt}(\ln p_{ijt}^{m}) - \hat{\mu}_{gt}(\ln p_{ijt}^{m*})}, \tag{15}$$

where $\hat{\mu}_{gt}(y_{ijt})$ denotes the estimate $\hat{\mu}_{gt}$ obtained using variable y_{ijt} as the outcome variable and recall that $p_{ijt}^{m} = (1 + \tau_{ijt}^{m})p_{ijt}^{m*}$ is the tariff-inclusive import price. If pre-tariff prices are unchanged, then the pass-through rate will equal one. Conversely, if pre-tariff prices

Table 2: Estimated Effects on U.S. Import Prices by Treatment Group and Period

Treatment Group	Pre-tariff Import Price			Tariff-inclusive		Pass-through	
	Pre-treat	2018	2019	2018	2019	2018	2019
Alum. and Steel 1	-0.011	***-0.112	***-0.104	**0.090	**0.100	0.45	0.49
	(0.022)	(0.031)	(0.039)	(0.037)	(0.041)		
Alum. and Steel 2	0.012	***-0.148	***-0.108	*0.052	0.016	0.26	0.13
	(0.019)	(0.026)	(0.026)	(0.030)	(0.033)		
China 1	-0.044	**0.173	0.070	***0.393	***0.276	1.79	1.34
	(0.031)	(0.068)	(0.078)	(0.068)	(0.079)		
China 2	-0.041	-0.004	**-0.102	***0.213	**0.103	0.98	0.50
	(0.035)	(0.042)	(0.052)	(0.042)	(0.052)		
China 3	-0.016	-0.040	**-0.105	0.049	0.047	0.55	0.31
	(0.032)	(0.040)	(0.050)	(0.040)	(0.050)		
China 4A	0.072		-0.053		0.048		0.47
	(0.045)		(0.049)		(0.051)		
E.U. Dispute	-0.027		***-0.299		*-0.122		-0.69
	(0.035)		(0.049)		(0.063)		
Solar and Washer	***0.417	0.047	*0.230	***0.283	***0.435	1.20	2.12
	(0.032)	(0.102)	(0.131)	(0.096)	(0.131)		
Weighted Average (excludes: Solar and Washer)							
All Goods	0.005	-0.027	**-0.082	***0.100	*0.064	0.79	0.44
	(0.021)	(0.031)	(0.038)	(0.031)	(0.038)		
Non-consumer Goods	-0.008	-0.024	**-0.081	***0.113	**0.077	0.83	0.49
	(0.020)	(0.030)	(0.038)	(0.029)	(0.039)		
Consumer Goods	0.029	-0.038	**-0.084	0.054	0.043	0.59	0.34
	(0.028)	(0.039)	(0.040)	(0.039)	(0.041)		

Notes: Pass-through rates are calculated by dividing the coefficient for tariff-inclusive prices by the difference in coefficients for tariff-inclusive and pre-tariff prices (see equation (15)). Standard errors clustered by product (8-digit) and by country in parentheses. $^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.

fall sufficiently such that tariff-inclusive prices are unchanged, then the pass-through rate will equal zero. The pass-through rate for retaliatory tariffs (ξ^x_{gt}) can be measured in the same manner. For imports, I estimate that pre-tariff prices fell by 0.118 and tariff-inclusive prices increased by 0.057, implying a pass-through rate of 0.33. For exports, I estimate that pre-tariff prices fell by 0.173 and tariff-inclusive prices also fell by -0.046, implying a pass-through rate of -0.36.

I can also assess the effects on U.S. prices by calendar time, which is more relevant for assessing the welfare impacts of the trade war. To do so, I begin by calculating the average ATT of each treatment group for the following three periods: pre-treatment, post-implementation in 2018, and post-implementation in 2019. The results for imports and exports are reported in Table 2 and 3, respectively. I postpone discussions of heterogeneity

Table 3: Estimated Effects on U.S. Export Prices by Treatment Group and Period

Treatment Group	Pre-tariff Export Price			Tariff-inclusive		Pass-through	
	Pre-treat	2018	2019	2018	2019	2018	2019
Alum. and Steel: Canada	-0.028	-0.042	-0.038	*0.060	***0.065	0.59	0.63
	(0.037)	(0.029)	(0.023)	(0.033)	(0.023)		
Alum. and Steel: China	-0.006	-0.077	***-0.318	***0.193	0.040	0.72	0.11
	(0.045)	(0.050)	(0.090)	(0.045)	(0.086)		
Alum. and Steel: Europe	0.022	-0.037	**-0.069	***0.143	***0.104	0.79	0.60
	(0.042)	(0.042)	(0.030)	(0.039)	(0.028)		
Alum. and Steel: India	0.020		*-0.048		***0.088		0.65
	(0.020)		(0.026)		(0.031)		
Alum. and Steel: Mexico	-0.004	***-0.045	0.003	*0.040	***0.086	0.47	1.04
	(0.016)	(0.015)	(0.014)	(0.021)	(0.021)		
China 1	0.016	-0.060	***-0.213	**0.087	-0.052	0.59	-0.32
	(0.038)	(0.045)	(0.071)	(0.044)	(0.071)		
China 2	-0.012	***0.191	-0.021	***0.398	***0.190	1.92	0.90
	(0.042)	(0.052)	(0.072)	(0.053)	(0.072)		
China 3	0.060	0.046	***-0.174	*0.106	-0.081	1.77	-0.87
	(0.037)	(0.055)	(0.055)	(0.055)	(0.056)		
China 4A	***0.254		***0.128		***0.152		6.45
	(0.066)		(0.048)		(0.049)		
Weighted Average (excludes: China 4A)							
All Goods	0.022	0.009	***-0.134	***0.125	-0.001	1.08	-0.01
	(0.026)	(0.033)	(0.045)	(0.033)	(0.046)		
Non-consumer Goods	0.028	0.031	***-0.132	***0.140	-0.003	1.28	-0.02
	(0.028)	(0.038)	(0.047)	(0.038)	(0.048)		
Consumer Goods	0.010	-0.039	***-0.137	***0.091	0.004	0.70	0.03
	(0.023)	(0.028)	(0.042)	(0.028)	(0.042)		

Notes: Pass-through rates are calculated by dividing the coefficient for tariff-inclusive prices by the difference in coefficients for tariff-inclusive and pre-tariff prices (see equation (15)). Standard errors clustered by product (6-digit) and by country in parentheses. $^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.

across treatment groups until Section 5 and merely note that the pre-treatment ATTs are statistically different from zero only for the Solar and Washer group for imports and China 4A for exports as previously mentioned. The ATTs for each period can then be averaged once more using respective treatment groups' trade values in 2017 as weights.

For 2018, I estimate that pre-tariff prices fell by 2.7% (without statistical significance) and tariff-inclusive prices increased by 9.5% with a pass-through rate of 0.79. For exports, the pre-tariff prices increased by 1% (without statistical significance) while tariff-inclusive prices increased by 13.3% with a pass-through rate of 1.08. While the 79% pass-through rate on import tariffs in 2018 is less than complete, it remains much higher than tariff pass-through estimates in other settings (see Fajgelbaum and Khandelwal (2022) for a review) or exchange rate pass-through, which Cavallo et al. (2021) found to be between 20-40%. Moreover, the pass-through rate is lower largely due to the lower tariff pass-through on aluminum and steel products (0.45 and 0.26), which is partly consistent with Amiti et al. (2020) who found a 0.5 pass-through rate for steel inputs. If we instead focus on the three waves of tariffs on China in 2018, the pass-through rate is 0.93, which is the same rate found by Cavallo et al. (2021) albeit using data covering 2019 as well. Thus, the staggered event study estimates for 2018 do not fundamentally change the conclusions of nearly complete tariff pass-through in earlier studies (Amiti et al., 2019; Fajgelbaum et al., 2020; Ma et al., 2021; Jiao et al., 2024) that focused on the initial impacts of the 2018–2019 trade war.

The estimates for 2019 show major departures from 2018. Most tariff-hit import varieties experienced statistically significant declines in their pre-tariff prices. As a result, tariff-inclusive import prices show statistically significant increases only for three groups: Alum. and Steel 1, China 1, and China 2. Overall, pre-tariff prices declined by 7.9% while tariff-inclusive prices increased by 6.6% with a pass-through rate of 0.44, a number more in line with estimates from other settings. Thus, only about 44% of the import tariffs were passed on to U.S. importers by 2019.

For export varieties, most treatment groups experienced statistically significant increases in tariff-inclusive prices despite lower pre-tariff prices. However, the two groups with the largest trade values (China 1 and 3) experienced particularly large drops in pre-tariff prices so that their tariff-inclusive prices actually declined although the effects are imprecisely estimated. In sum, I estimate that pre-tariff prices declined by 12.5% while tariff-inclusive prices did not change at all, which implies a zero pass-through rate. U.S. exporters on average ended up absorbing all of the retaliatory tariffs in 2019.

4.3 Welfare Effects

Using these estimates, I return to the equivalent variation in equation (3) to assess the impacts of the 2018–2019 trade war tariffs on U.S. welfare. The equation is reproduced here for convenience.

$$EV_t^0 = \underbrace{\sum_{g \in G^M} R_{g0} \cdot ATT_{gt}^{qm}}_{\equiv QM_t^0} + \underbrace{\left(- \sum_{g \in G^M} M_{g0} \cdot ATT_{gt}^{pm*} \right)}_{\equiv PM_t^0} + \underbrace{\sum_{g \in G^X} X_{g0} \cdot ATT_{gt}^{px*}}_{\equiv PX_t^0}$$

The ATTs in the second and third terms of the equation (i.e., pre-tariff import and export prices) have already been estimated and discussed in the previous sections. The ATTs in the first term requires estimating the staggered event study specification in equation (14) using log import quantity as the outcome variable and weighting by initial tariff revenues. I discuss the results of this estimation in Section A.3. Once all the ATTs in the equation have been estimated, I only need to choose the appropriate unit of time for the calculation. I first calculate the average ATT for each calendar year and treatment group and use them to calculate the equivalent variation for year 2018 and 2019. Table 4 presents the calculated equivalent variations and also expresses them as percentages of U.S. GDP in 2017. The three decomposed terms in equation (3) are also reported.

Table 4: Impacts of the 2018–2019 Trade War on U.S. Welfare

Year	Unit	Welfare Effect (EV_t^0)	Tariff Revenue (QM_t^0)	Import Price (PM_t^0)	Export Price (PX_t^0)
2018	% of GDP	0.01	0.00	0.01	-0.00
	Billion $	1.56	0.56	1.49	-0.49
2019	% of GDP	0.05	-0.00	0.14	-0.08
	Billion $	9.89	-0.75	26.76	-16.12

Notes: Percentage of GDP calculated using U.S. GDP in 2017.

In line with the results in the previous section, both the overall and decomposed impacts on U.S. welfare were negligible in 2018. In contrast, I estimate a 0.05% (or $9.89 billion) *gain* in U.S. welfare for 2019. The decomposition further shows that this was driven entirely by a net terms-of-trade gain. The tariff revenue effect from decreased quantity of imports remains negligible due to the fact that U.S. import tariff were close to zero in 2017 and collected relatively small amounts of tariff revenue. Despite larger decreases in the pre-tariff prices of exports, the net terms-of-trade gain turns out as positive due to the fact that the 2018–2019

trade war impacts a much larger value of U.S. imports than exports. This welfare impact is similar in magnitude to the welfare losses estimated by Amiti et al. (2019) and Fajgelbaum et al. (2020) for 2018 ($8.2 billion and $7.2 billion, respectively) but has the opposite sign.

5 Implications and Sources of Heterogeneous Price Responses

5.1 Implications for Prices of Consumer and Non-consumer Goods

The aggregate effects discussed in Section 4.2 mask substantial heterogeneity across treatment groups that have important implications for U.S. prices. As the summary statistics presented in Table 1 show, the earlier import tariff waves almost entirely landed on non-consumer goods. It was only with the third tariff wave against China that consumer goods started taking up a significant share of targeted imports. Interestingly, Table 2 shows that the tariff-inclusive prices of imports hit by earlier tariff waves are the ones that experienced statistically significant increases by 2019. For the later tariff waves that affected consumer goods, the increase in tariff-inclusive prices were more muted and not statistically different from zero. Thus, U.S. importers of consumer goods would not have felt tariff-induced price increases as strongly as those importing non-consumer goods.

To see this more clearly, I take the trade-weighted average across treatment groups once again. However, I use each treatment group's 2017 import values of consumer and non-consumer goods as weights instead of the import value of all goods used in Section 4.2. For example, Table 1 shows that group China 3 had an import value of about $187 billion but consumer goods only took up 32% of this import value. Thus, the import value of consumer goods for this group will be about $60 billion and the remainder goes to non-consumer goods. The last two rows of Table 2 report the weighted averages. I find that pre-tariff prices fell by about 8% for both non-consumer and consumer goods in 2019, but final import prices saw statistically significant increases of about 8% only for non-consumer goods. The estimate for consumer goods is about half at 4% and not statistically significant. As a result, the pass-through rate for non-consumer goods is also higher at 0.49 and only 0.34 for consumer goods. In comparison, the same calculation for export prices in Table 3 reveals no difference across consumer and non-consumer goods.

Cavallo et al. (2021) find that U.S. retail prices did not show significant increases as a result of the trade war. Given that they found nearly complete pass-through rates for import tariffs, they logically concluded that U.S. retailers absorbed much of the higher import prices. The relatively muted effects on the import prices of consumer goods provides

a complementary explanation on why U.S. retail prices did not experience visible increases during 2018–2019.

5.2 Import Market Share and Exporter Price Responses

What could explain the heterogeneity in price responses? The theory of international trade suggests that tariffs imposed by large importers can induce foreign exporters to lower the price of their goods. Intuitively, the idea is that when the importer's demand takes up a large share of the total demand for the foreign exporter's output, the tariff-induced reduction in import demand will be significant enough to shift the demand curve to the left. As long as the foreign exporter's supply curve is upward sloping, this leftward shift in demand will lead to a lower equilibrium pre-tariff price of exports.

I empirically test this idea by measuring the U.S. and its trade partners' market share in the exporting country's export market for each 6-digit product. To explain this process, I first need to enrich the notation to distinguish importers from exporters. Let m_{cij0} denote importing country c's import value of product i from exporting country j in 2017. Similarly, let x_{cij0} denote exporting country c's export value of product i to importing country j in 2017. Then, a country c's importer market share for product i in the exporting country j is measured by

$$\lambda^m_{cij0} = \frac{m_{cij0}}{\sum_k x_{jik0}}. \tag{16}$$

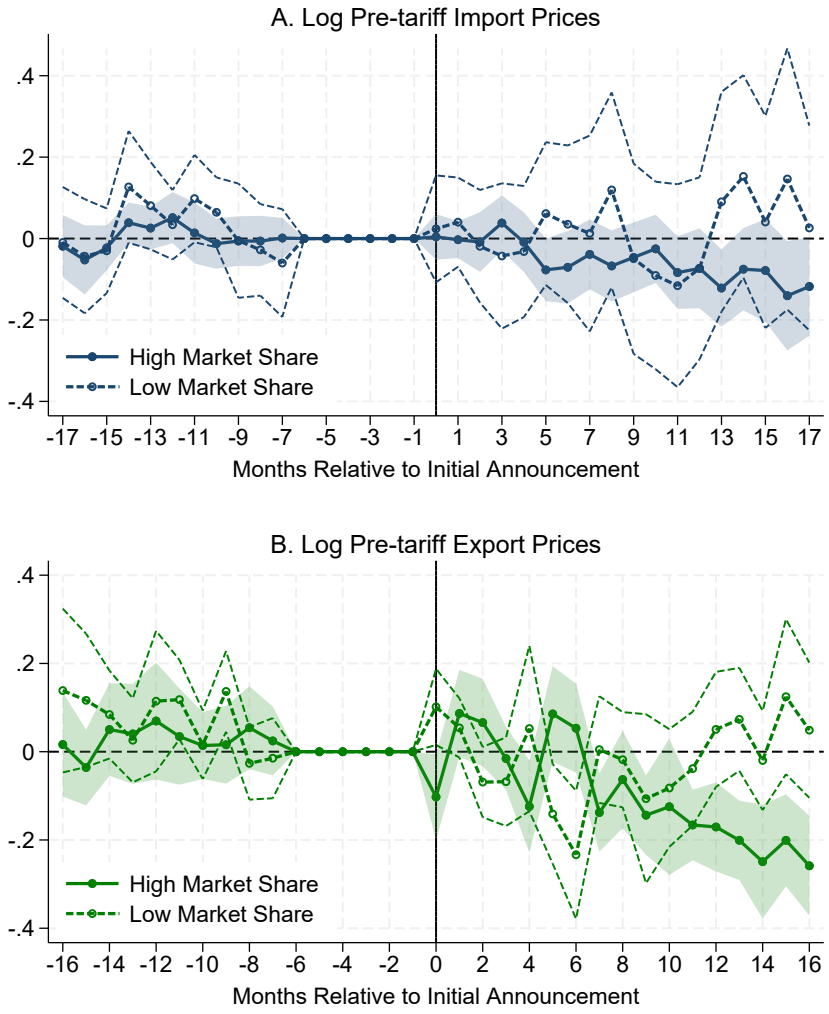
For example, a share of $\lambda^m_{cij0} = 1$ indicates that importer c is the sole source of demand for exports of product i produced in country j. In this case, we can expect changes in country c's import demand to have a significant impact on country j's exporters of product i.

I use international trade data from WITS at the 6-digit product level to measure importer market shares in equation (16) for the U.S. and its trade partners. Within treated export varieties, I find that the retaliating country for the median variety had an importer market share of 3.7% with a standard deviation of 14 percentage points. In comparison, within treated import varieties, the U.S. had an importer market share of 14.6% for the median variety with a standard deviation of 20 percentage points. Therefore, we can see that the U.S. held a much larger share of foreign exporters' markets it hit than the retaliating countries did in the U.S. exporters' markets.

To further investigate whether the heterogeneity in price responses vary along importer's market share, I first partition the sample of varieties into two: those with high and low importer market shares. The threshold I use for this partition is the 3.7%, which is the median value calculated from treated export varieties. Then, I estimate the staggered event

study specification in equation (14) using the partitioned samples for import and export varieties. The trade-weighted ATTs are plotted in Panel A and B of Figure 3, which are comparable to the plots in Panel A and B of Figure 2.

Figure 3: Estimated Effects on U.S. Import and Export Prices by Importer's Market Share



Notes: Enclosed regions around point estimates are 95% confidence intervals based on standard errors clustered by product (8-digit for imports and 6-digit for exports) and by country. Import and export varieties are in the "High Market Share" sample if and only if the importing country's share of the exporting country's total export market for the associated 6-digit product was more than 3.7%. See equation (16) for more details.

For both imports and exports, the declining trend in pre-tariff prices previously observed in Figure 2 are only visible among varieties where the importer held high market shares. For other varieties, pre-tariff prices remained unchanged and statistically indistinguishable from zero. The contrast is more precisely estimated for export varieties in line with the main finding that the pre-tariff prices fell more significantly for exports. These empirical patterns lend support to trade theory's prediction that large importers are able to use tariffs

to manipulate their terms of trade.

I end this section by noting that importer size plays an independent role from import demand or export supply elasticities, which play important roles in determining the optimal tariff rates (Broda et al., 2008). For instance, Cavallo et al. (2021) find that pre-tariff export prices fell primarily for non-differentiated products and reason that it would have been easier for importers to find substitutes if prices remained the same. Even with essentially homogeneous goods, however, the exporters will not be willing to lower their prices in response to an importing country's tariffs if only a negligible share of their revenue is derived from the importer. Thus, the evidence presented in this section complements the analysis in Cavallo et al. (2021) as opposed to being redundant.

6 Conclusion

In this paper, I reassessed the 2018–2019 trade war's staggered impacts on U.S. prices and welfare by taking advantage of recent econometric advancements in the DiD literature. I found that the pooled event study produces unbiased estimates for import values but not import prices. Relying on results from Sun and Abraham (2021), I was able to trace this discrepancy to the fact that the tariffs' effects were heterogeneous for import prices but not import values. Adapting the estimation procedure proposed in Sun and Abraham (2021) to the current setting, I find evidence that both U.S. and foreign exporters eventually lowered their prices in response to the trade war tariffs by 2019. Thus, I also estimate lower tariff pass-through rates for both import and retaliatory tariffs compared to previous studies.

Although pre-tariff prices were lowered more significantly for U.S. exports than imports, the 2018–2019 affected a much larger value of U.S. imports. As a result, the U.S. experienced a small welfare gain stemming from a positive net terms-of-trade effect. Nevertheless, the positive welfare gain should not be interpreted as a confirmation of U.S. trade policy's ultimate success. Rather, it is an empirical confirmation of trade theory's prediction that import tariffs levied by large importers can induce foreign exporters to lower their prices, resulting in terms-of-trade gains at least in the short run.

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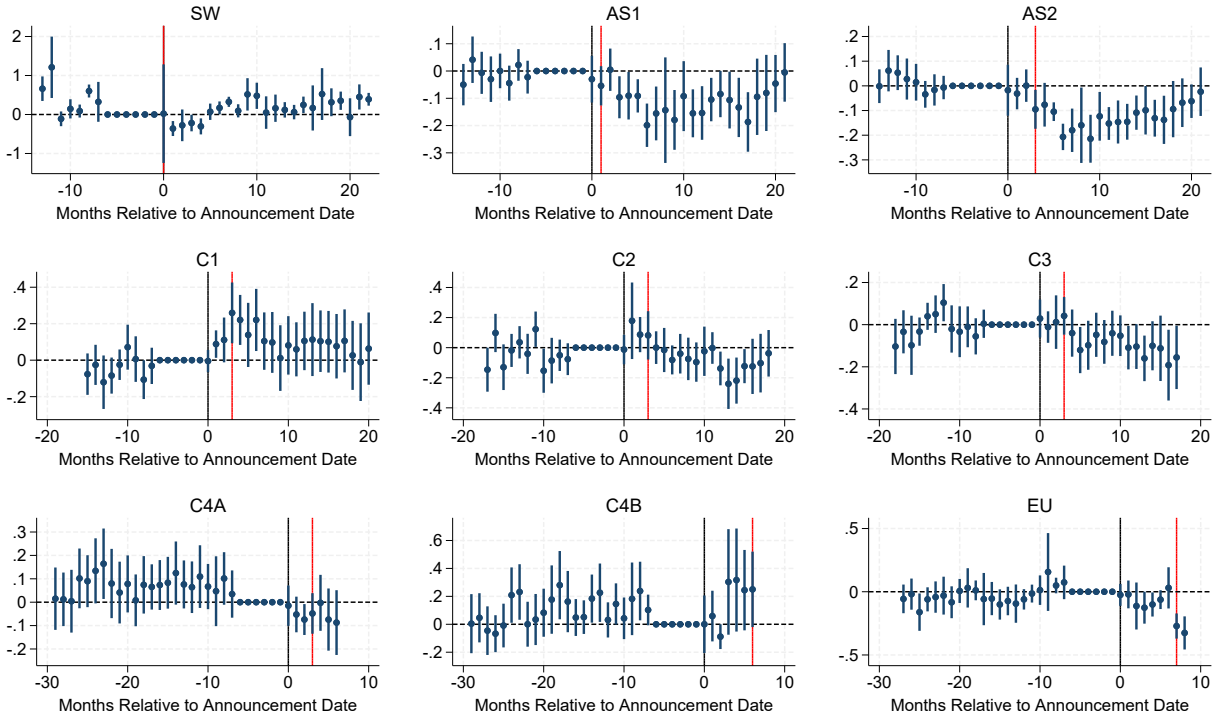
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A Appendix

A.1 Event Study Plots for Pre-tariff Prices and Values of Imports

Figure 4 and 5 document the staggered event study estimates for pre-tariff import price and import value by treatment group.

Figure 4: Individual Staggered Event Study Estimates for Pre-tariff Import Prices



Notes: The red vertical line marks the tariff implementation date. The bars around point estimates are 95% confidence intervals based on standard errors clustered by product (8-digit) and by country.

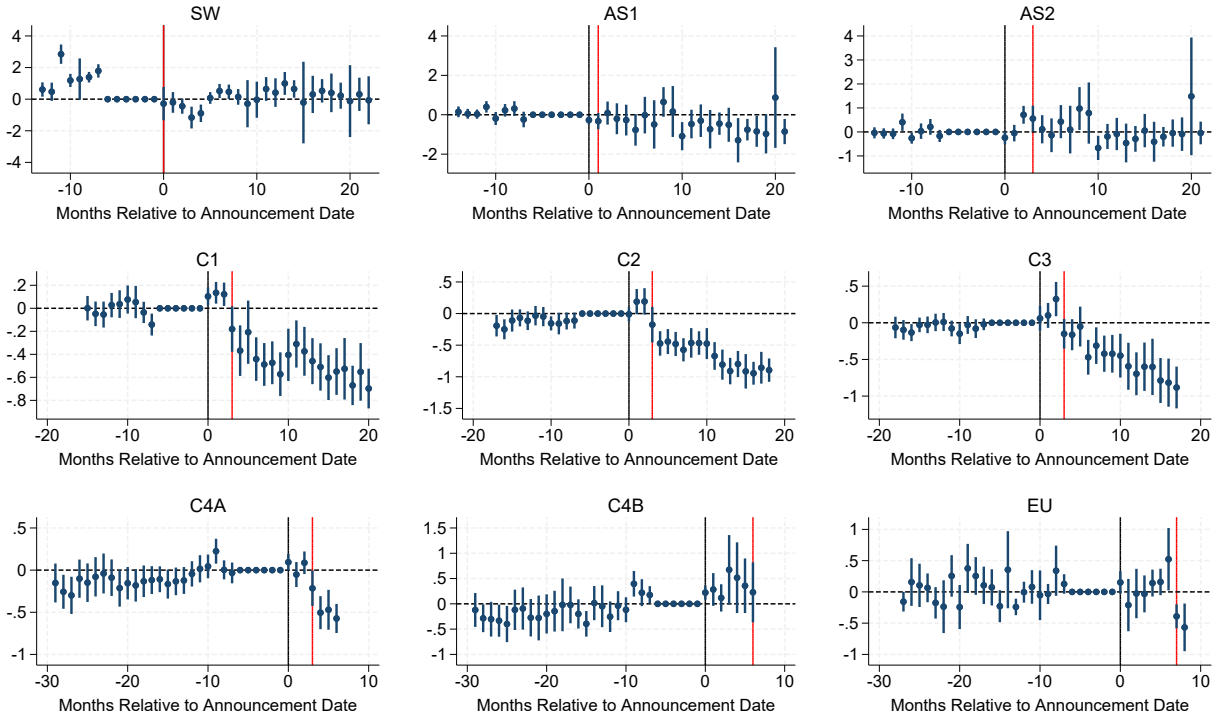
A.2 Event Study Plots for Pre-tariff Prices and Values of Exports

Figure 6 and 7 document the staggered event study estimates for pre-tariff import price and import value by treatment group.

A.3 Event Study Plots for Import Quantities

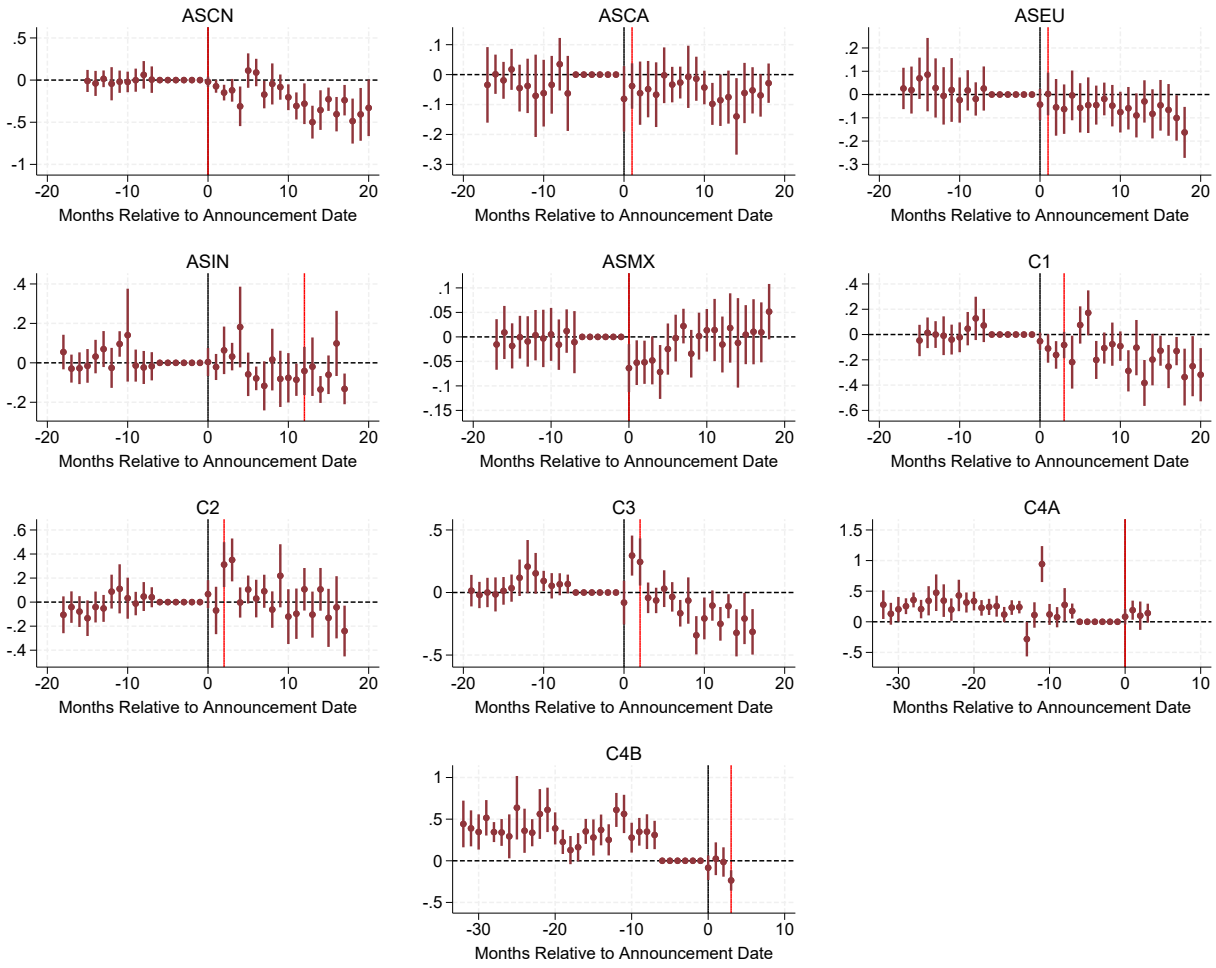
Figure 8 documents the staggered event study estimates for import quantities. Note that the estimation here weights observations by their initial tariff revenues as opposed to import values in accordance with equation (3).

Figure 5: Individual Staggered Event Study Estimates for Import Values



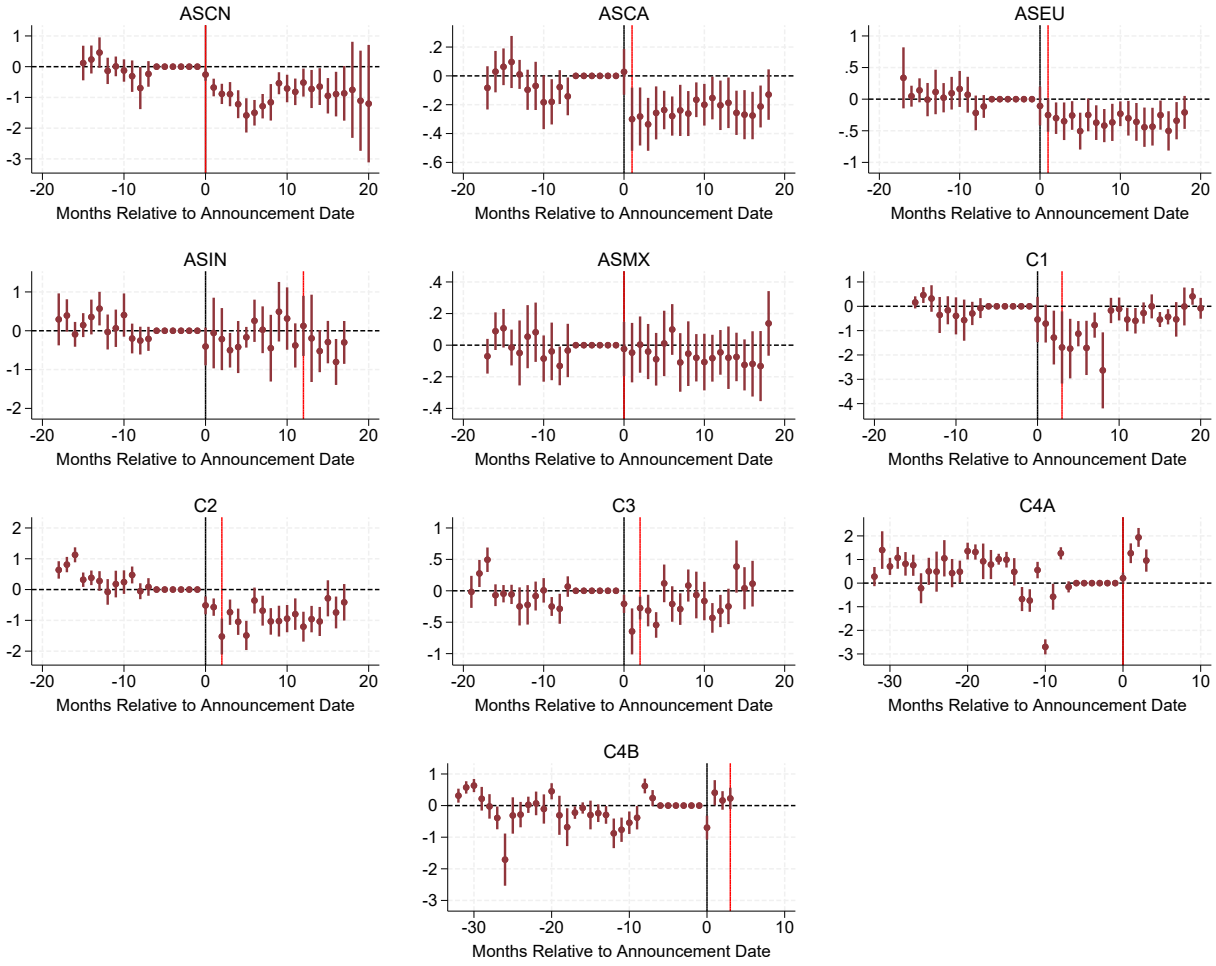
Notes: The red vertical line marks the tariff implementation date. The bars around point estimates are 95% confidence intervals based on standard errors clustered by product (8-digit) and by country.

Figure 6: Individual Staggered Event Study Estimates for Pre-tariff Export Prices



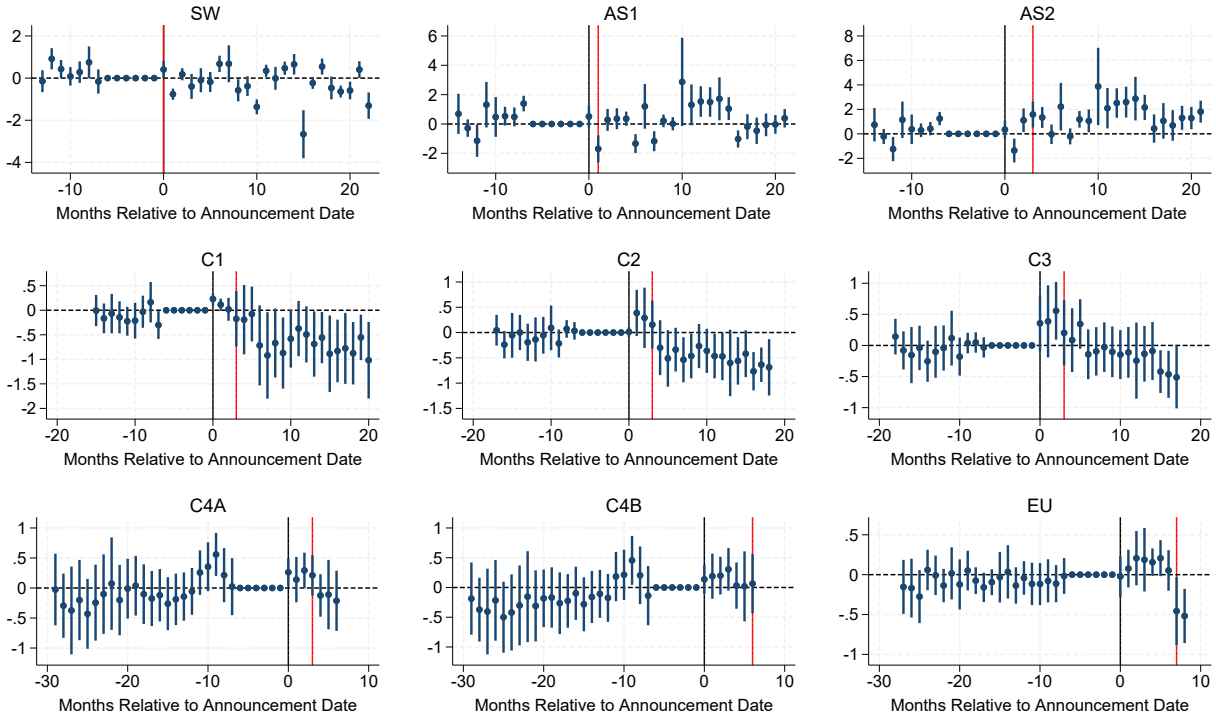
Notes: The red vertical line marks the tariff implementation date. The bars around point estimates are 95% confidence intervals based on standard errors clustered by product (6-digit) and by country.

Figure 7: Individual Staggered Event Study Estimates for Export Values



Notes: The red vertical line marks the tariff implementation date. The bars around point estimates are 95% confidence intervals based on standard errors clustered by product (6-digit) and by country.

Figure 8: Individual Staggered Event Study Estimates for Import Quantities



Notes: The red vertical line marks the tariff implementation date. The bars around point estimates are 95% confidence intervals based on standard errors clustered by product (8-digit) and by country.