Tariff Pass-through into Retail Prices: Measurement, Replacement, and Shrink-flation*

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Abstract

To what extent do import tariffs pass-through into retail prices? Existing estimates for the U.S.-China trade war present a puzzle: pass-through into U.S. border prices was complete with seemingly no effect on retail prices. We combine barcode country-of-origin data with retail scanner data to address this apparent discrepancy in the context of consumer goods. Within barcodes, the tariff elasticity of retail prices is 0.14, yet the tariff elasticity of aggregated prices per unit weight, a measure similar to the unit values recovered from customs data, is 0.44. This discrepancy is attributable to product replacement bias: pass-through occurs predominantly via product turn-over towards barcodes with smaller package size, even within narrow firm-product combinations. When accounting for the foreign cost share of U.S. retail imports, implied tariff pass-through into retail prices ranges from 60%-100%.

JEL Categories: E31, F13, F14, F31, L11.

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1 Introduction

Who pays for import tariffs? Basic economic theory suggests that the rate of tariff pass-through - the extent to which the buyer price reflects a changing tariff - is the key metric for quantifying tariff incidence across foreign producers and domestic consumers. Empirical estimates of tariff pass-through, however, are scarce, given the remarkable liberalization of global trade policy in the post-war period.

The U.S.-China trade war of 2018-2019 therefore offers a unique setting in which to study both the outcomes of this particular policy episode, but also the broader incidence of tariffs across importers and exporters, foreign producers and domestic consumers. Almost all research in this vein convincingly points to the same finding: U.S. tariffs applied to Chinese exports in 2018-2019 were passed through completely into prices paid by U.S. importers, such that U.S. buyers bore the brunt of tariffs (Amiti et al. 2019; Fajgelbaum et al. 2020; Cavallo et al. 2021; Gopinath and Neiman 2026; Hinz et al. 2026). And yet there remains remarkably little evidence for any tariff pass-through into prices faced by U.S. households at U.S. stores.[1]

In this paper we use barcode-level scanner data to estimate the effect of these tariffs on final consumer prices. We adopt a broad definition of pass-through that encompasses not only tariff-induced price changes for existing barcodes, but also the extent to which tariffs might generate higher prices via product replacement. For every 10% increase in tariffs, retail prices (per unit weight) increase by 4.4%. To the best of our knowledge, this is the first paper to document price effects of this magnitude for consumer goods during the U.S.-China trade war, and to show that product replacement was the predominant mechanism through which tariffs led to higher prices. We show that our estimate of the tariff elasticity of retail prices is compatible with complete pass-through in price levels, despite being incomplete in terms of percentage changes. Since tariffs are only applied to import values recorded at the border, a low foreign cost share of retail imports generates muted estimates of tariff pass-through into retail prices. By estimating this foreign cost share directly, we find that tariff pass-through

[1] Cavallo et al. (2021) estimate pass-through into U.S. retailer prices of only 3.5%, and this estimate is not distinguishable from zero at standard confidence levels. See Fajgelbaum and Khandelwal (2022) for more discussion.

into retail prices was between 60%-100%, such that the incidence of import tariffs fell largely on U.S. consumers.[2]

We combine several datasets to estimate the response of consumer prices to tariffs. Our primary source is NielsenIQ data on barcode-level prices and sales across approximately 70 retail chains. We merge these data with firm identifiers and barcode-level country-of-origin information and use a crosswalk between NielsenIQ product modules and HS8 tariff codes to link products to the 2018/19 waves of tariffs on Chinese imports.

We begin by estimating the retail price elasticity of tariffs measured using unit values aggregated to the firm, category, production country, and retailer level. For example, the average price per unit weight of a specific firm's barcodes of deodorant produced in China and sold at a specific retail chain in the U.S. By implementing a triple-difference design comparing changes in unit values within production origins and within product categories, we recover a headline unit value tariff elasticity estimate of 0.44.

While this estimate is considerably larger in magnitude than existing tariff elasticity estimates associated with U.S. retail prices, it still appears to fall short of the full pass-through consistently recovered using customs data. A distinguishing feature of the dataset used in this paper is that we observe both price and quantity data for detailed product definitions – barcodes – over time, as well as country of origin information at the same level of disaggregation. The combination of consistent product definitions with price and quantity sales data allows us to compare estimates of tariff elasticities at both the barcode level and at higher levels of aggregation, such as unit values within a given firm and product category. We leverage this novel feature of the data, and find that our retail unit value tariff elasticity of 0.44 is in fact entirely consistent with full, or near-full, pass-through into border prices, as well as the existing null estimates of pass-through into retail prices.

First, we show that the tariff elasticity of retail unit values may be smaller in magnitude than similar percentage pass-through estimates at the border despite reflecting the exact same magnitude of pass-through in levels. In the presence of some cost wedge associated with post-tariff distribution, such as retailing mark-ups, retailing costs and/or intra-national

[2]Throughout this analysis, we focus exclusively on tariff waves 3 and 4a associated with the U.S.-China trade war. Waves 1 and 2 primarily targeted goods with poor representation in scanner data, such as intermediates and capital goods. Section 2 provides more detail.

distribution costs, estimates of pass-through into retail prices must be scaled by the value share of retail imports that is attributable to pre-tariff activity—the foreign cost share—in order to be comparable to estimates of pass-through at the border.[3]

Despite the crucial role played by the foreign cost share in shaping pass-through of international shocks to retail prices, estimates are scarce. Existing estimates tend to place the foreign cost share for U.S. retail between 30%-50% (Berger et al. 2012; Flaaen et al. 2025) which overlaps almost exactly with the confidence interval of our headline estimate of the tariff elasticity of retail unit values. However, the product categories reported in (Berger et al. 2012) map only imperfectly into a subset of our product categories and are based on older data (1994-2007). We therefore carry out our own estimation of the foreign cost share, by mapping U.S. customs unit values to scanner data unit values across product categories and origin countries, with an average recovered estimate of 32% for consumer retail goods imported from China.[4] We provide a qualitative validation of this estimate by documenting a positive relationship between our foreign cost share estimates and levels of tariff pass-through into retail unit values, and recover a pass-through estimate in the range of 60%-100%. To the best of our knowledge, this is the first paper to document the importance of the foreign cost share for measuring tariff pass-through into final consumer prices for a broad set of consumer good categories.

Second, we show that while our headline estimates align with existing estimates of full pass-through into border prices, they are not incompatible with the null estimate of product-specific tariff pass-through reported by Cavallo et al. (2021). Specifically, we find that the retail unit value tariff elasticity is only 0.10 when we restrict the sample to continuing barcodes.[5] Similarly, when we estimate the tariff elasticity using within-barcode variation in a balanced panel of barcodes, we recover an estimate of 0.14.[6] This within-barcode

[3]For example, if the cost share of domestic economic activity associated with retail goods, such as intra-national distribution costs and retailing costs (land, labor, etc.), is 60%, a tariff is only applied to 40% of the value of the good. Full pass-through at the border would imply 40% pass-through into retail prices. See Sangani (2025) for a similar argument regarding commodity shock pass-through.

[4]In contemporaneous work, Flaaen et al. (2025) also consider the importance of the foreign cost share for tariff pass-through in their study of EU wine exports to the U.S. They estimate a foreign cost share for EU wine that is similar to our estimate, at 38%.

[5]We define a continuing barcode as one with non-zero sales throughout the entirety of our sample period.

[6]We confirm that our relatively low estimates of within-barcode pass-through are not driven by endogenous price spillover effects to non-tariff goods, tariff "smoothing" by retailers, or the effect of tariff shocks

estimate is reduced to zero once we expand our sample to an unbalanced panel of barcodes, similar to Cavallo et al. (2021). This suggests that price increases for the same barcodes are only introduced for products that continue to be sold for a sufficiently long time after the introduction of tariffs.

A significant share of the tariff elasticity of retail prices therefore occurs via product replacement, such that within-barcode estimates differ considerably from unit value estimates. We decompose our unit value tariff elasticity estimate into changes in the numerator—what we call "sticker prices"—and the denominator, or average package size.[7] We recover a tariff elasticity of average package size of -0.29, accounting for approximately two thirds of our unit value tariff elasticity. We refer to this effect as "shrink-flation", and confirm a lack of pre-trends, suggesting this decrease in size is causally attributable to the tariff change. Our findings therefore suggest that within-product measures of pass-through are potentially subject to product replacement bias (Nakamura and Steinsson 2012), such that within-barcode estimates may under-state the true pass-through of import tariffs into consumer prices.

Still, within-barcode estimates have the benefit of a clear counterfactual price, whereas entry and exit could lead to compositional shifts in the set of products offered, thereby generating price elasticity estimates which are not necessarily relevant for consumer welfare.[8] There are several reasons to think that the unit value tariff elasticity estimates we recover are likely welfare relevant, however. First, we estimate these effects within narrow firm-country-category-retailer cells, such that changes in average package size occur within fairly narrow product definitions. Second, it would have to be that quality is almost perfectly correlated with the inverse of package size in order for our recovered tariff elasticity of package size estimates to be driven by an endogenous quality response, rather than product replacement. Finally, we document a positive tariff elasticity of barcode entry, with tariff-induced entry concentrated in barcodes with small package size. In contrast, the effect of tariffs on barcode exit is more muted and not concentrated in any portion of the barcode size distribution.

raising all production costs and thereby reducing the effect of product-specific tariffs.

[7] By "sticker price", we refer to the average price per barcode sold, rather than the average price per unit of weight sold, which we refer to as unit values. Section 4 provides an exact decomposition of unit value pass-through into a sticker price component and an average package size component.

[8] As discussed in Metzler (1949) and documented by Ludema and Yu (2016); Chen and Juvenal (2016); Freitag and Lein (2023), tariffs might generate an endogenous exit of low-quality varieties, thereby mechanically increasing unit values even in the absence of any within-variety pass-through.

Finally, we consider the role of buyer and seller power in shaping heterogeneity in observed tariff elasticities across products and retailers. Specifically, we interact tariff changes with both the share of expenditure within a given retailer attributable to a specific firm (the seller share) and the share of expenditure within a given firm attributable to a specific retailer (the buyer share). In line with notions of monopoly and monopsony power, we find elevated tariff price effects for above-median seller shares and reduced tariff price effects for above-median buyer shares. This heterogeneity holds for both unit values and barcode-specific prices. Yet firms with above-median seller shares exhibit a pronounced gap between unit value elasticities and barcode-specific price elasticities, suggesting that endogenous product replacement is particularly common when seller power is elevated. Alternatively, when firms exhibit below-median seller power at below-median retailers, barcode price and unit value tariff elasticity estimates are effectively identical, at 0.17. These heterogeneous tariff elasticities suggest that product replacement may be important for both measuring aggregate effects of cost shocks and measuring the heterogeneous effects of such shocks across producers and retailers.

Related literature: This paper fits into a growing empirical literature on the economic consequences of the U.S.-China trade war, with a particular focus on the price effects of this event (Amiti et al. 2019; Fajgelbaum et al. 2020; Flaaen et al. 2020; Cavallo et al. 2021).[9] The closest papers to ours are Cavallo et al. (2021) and Flaaen et al. (2025) who study pass-through in the context of, respectively, consumer retail goods and E.U. wine exports. Our null estimate of within-barcode pass-through for an unbalanced panel matches findings in Cavallo et al. (2021), who consider a similar specification. By estimating pass-through into retail unit values, we show that product replacement can align this result with full pass-through at the border. Similar to Flaaen et al. (2025), we show that domestic wedges associated with inter-mediation of imported goods generate muted price effects in retail data that are not necessarily inconsistent with full pass-through at the border.[10]

[9]See Fajgelbaum and Khandelwal (2022) for a full literature review of the first U.S.-China trade war. Key papers include Charbonneau and Landry (2018); Fetzer and Schwarz (2021); Benguria et al. (2022); Blanchard et al. (2024); Handley et al. (2025); Hankins et al. (2025). See Bown (2021) for a full timeline of the events which constitute the U.S.-China trade war. See Cavallo et al. (2025) and Amiti et al. (2026) for estimates of tariff pass-through following the "Liberation Day" tariffs announced on April 1, 2025.

[10]Ganapati and Hottman (2026) argue that pass-through to U.S. importers was only 60% despite an increase in import prices consistent with full pass-through: scale economies and a downward-sloping export

Our results build on the product replacement bias documented in Nakamura and Steinsson (2012).[11] Producers and/or retailers may perceive product replacement as a mechanism for increasing prices which generates less negative sentiment among consumers (Rotemberg 2005; Nakamura and Steinsson 2011), and a growing literature suggests that "shrink-flation" may be a particularly common method by which firms achieve higher prices when replacing products (Evangelidis 2024; Chalioti and Serfes 2024). We similarly contribute to the literature quantifying the foreign cost share of domestic retail goods (Burstein et al. 2003; Devereux et al. 2003; Anderson and Van Wincoop 2004; Goldberg and Campa 2010; Berger et al. 2012), pass-through of exchange rate shocks to retail prices (Auer et al. 2021), and the combination of pass-through estimates along the distribution network (Nakamura and Zerom 2010; Koujianou Goldberg and Hellerstein 2013; Flaaen et al. 2025). By studying the endogenous response of retailers to tariffs, we also contribute to the literature documenting the relationships between foreign producers, domestic retailers, and trade shocks (Raff and Schmitt 2006, 2009; Eckel 2009; Basker and Van 2010; Basker et al. 2012; Raff and Schmitt 2012, 2016; Cole and Eckel 2018; Jaccard 2023).

Finally, we provide novel empirical evidence regarding pass-through by seller shares and buyer shares within the context of consumer retail goods (Atkeson and Burstein 2008; Amiti et al. 2016; Alviarez et al. 2023). We find that larger producers pass-through a higher share of the tariff cost shock, while larger retailers reduce the level of pass-through to consumer prices. Furthermore, the product replacement mechanism we document appears to be most relevant for larger producers, because the gap between unit value and product-specific price pass-through is largest for firms with some monopolistic power.

Section 2 describes our data. Section 3 provides our headline estimates of pass-through into retail unit values alongside estimates of the foreign cost share. Section 4 decomposes unit value pass-through into components associated with entering, exiting, and continuing barcodes, and Section 5 concludes.

supply curve generate price effects which exceed the incidence of import tariffs. From the perspective of retail prices, the focus of our paper, the cost shock remains equivalent to full tariff pass-through at the border, and we therefore maintain full pass-through at the border as our benchmark.

[11] While product entry is beneficial to consumers (Feenstra 1994; Broda and Weinstein 2006), product replacement bias suggests not all entry is equal. If product entry serves only as a means to increasing the price of an otherwise identical product, such "entry" should not be considered relevant for consumer welfare.

2 Data

We bring together several data sources for the empirical analysis of this paper: i) barcode-level price and expenditure from NielsenIQ Retail Scanner Data, ii) barcode-level country-of-origin data from both Label Insight and scraped from the website of a large U.S. retailer, iii) firm identifiers from GS1, iv) tariff data and import unit values at the eight-digit Harmonized System (HS) level from the U.S. Trade Representative (USTR), and v) a concordance between NielsenIQ product modules and HS-10 digit codes (constructed by the authors).[12]

2.1 Price and Expenditure

For information on prices and expenditures at the barcode-level, we make use of the NielsenIQ Retail Scanner Panel over the period 2015-20. Products from all NielsenIQ departments are included in the data, such as alcoholic beverages, dairy, deli, dry grocery, fresh produce, frozen foods, packaged meat, general merchandise, health and beauty care, and non-food grocery. Depending on the year, data are included from approximately 30,000-50,000 participating grocery, drug, mass merchandiser, and other stores.

The raw data is available at the barcode-store-week level. For the analysis, we build a panel dataset of prices and expenditures at the barcode-retailer-month level. To abstract away from changes in the set of reporting stores over time, we focus on the set of stores that report positive sales in every year of our sample period ($n = 29,054$). To aggregate sales and prices at the retailer level, we make use of the NielsenIQ retail chain identifiers.

2.2 Country of Origin

One limitation of the Retail Scanner Data when used in empirical studies of international trade is the lack of information regarding a product's country of origin. To overcome this, we supplement scanner data with two separate data sources that provide origin country information at the barcode level. First, we use a proprietary dataset purchased from Label Insight, a private firm specializing in the provision of detailed product attribute data. Second,

[12] This is an extension of the concordance between NielsenIQ product modules and HS-6 digit codes used in Bai and Stumpner (2019).

we scrape data from the website of a large U.S. retailer, which further improves our product coverage.

2.3 Tariff Data

We use data from the official tariff lists published by the United States Trade Representative (USTR) for waves 1-4. Figure 1 shows the weighted average tariff that the U.S. Government applied to imports from China between 2017-2021, using 2017 import volumes as weights. While the average tariff increase applied to NielsenIQ consumer product modules was significantly lower than for other targeted categories, the tariff increase applied to Chinese goods within the NielsenIQ data from early 2017 to the peak at the end of 2019 was still over 7ppt, compared to a pre-2018 baseline average tariff rate of around 3%.[13]

There were four waves of tariff changes associated with the 2018-2019 trade war, each implemented separately: Waves 1, 2, 3, and 4a. Figure 2 plots the tariff rate applied to Chinese goods by wave, as well as the average tariff increase applied to all imports from China. The targeted categories account for roughly two thirds of all US imports from China in 2017, while the remaining third (list 4b) were left as a threat point during the U.S.-China negotiations and were eventually left untaxed after the signing of the Phase One Deal in January 2020.

In the main part of our analysis we focus on the comparison between Wave 3 and the never taxed categories. Waves 1 and 2 focused largely on intermediate goods and have a poor representation in NielsenIQ scanner data. While our scanner data contains a sufficient number of products across waves 3 and 4a to include both in our analysis, we focus our results on those associated with wave 3, with wave 4a results provided as a robustness. We do this for two reasons. First, as shown in Figure 2, the wave 3 tariffs experienced two increases: an increase from 2.5% to 12.5% in September 2018, and a further increase to 27.5% in May 2019. Wave 3 goods then maintained this heightened tariff rate through 2020. Wave 4a categories, on the other hand, experienced a tariff increase from 5% to 20% in August 2019 which was followed by a reduction to 12.5% in early 2020. The dynamic effects associated

[13]The initial two waves of the U.S.-China trade war focused heavily on intermediates (chemicals and machinery), rather than consumer goods.

Figure 1: Trade War Tariffs by Dataset

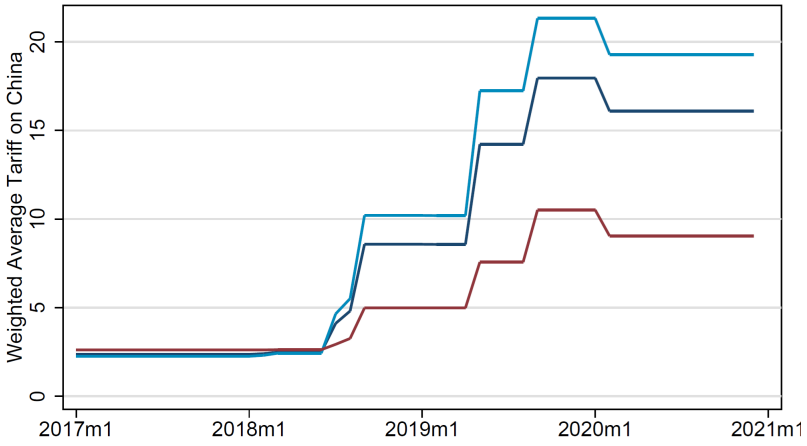
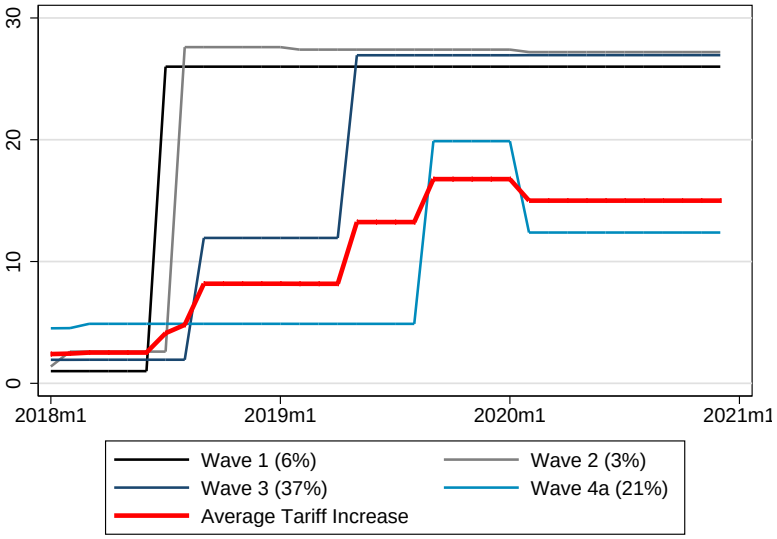


Figure 2: Trade War Tariffs by Wave within NielsenIQ Sample



Notes: Figure 1 provides the average tariff rate applied by the U.S. to imports from China over time. All products are in dark blue, products outside of the NielsenIQ scanner data are in light blue, and products within NielsenIQ scanner data in dark red. Figure 2 provides the average tariff applied to Chinese imports over time within NielsenIQ categories, and denotes the total share of categories within the NielsenIQ data associated with each tariff wave during the 2018-2019 trade war. These shares do not sum to one as the remaining NielsenIQ categories had not tariff change, and therefore constitute a control group.

with a short-lived, dramatic, increase in tariff rates followed by a mediation of rates to 12.5% complicates effect interpretation, particularly given the lagged effects of tariffs we document below. Second, since wave 4a tariff rates stabilize at the start of 2020, there is only a short period prior to the Covid-19 pandemic with which to estimate tariff effects. While we control for country-month and product category-month shocks, as well as seasonality controls specific

to tariff-affected goods, it is still possible that the Covid-19 pandemic induced a distribution of demand and supply shocks that are unobserved and potentially correlated with treatment, which we cannot account for here. Still, we show below that at least qualitatively, and to a large extent in magnitudes as well, estimates associated with the wave 4a tariffs had similar effects to the wave 3 tariffs.

2.4 Concordance between NielsenIQ and tariff data

Barcode price and expenditure data in NielsenIQ are organized in roughly 1,200 different product modules. In order to apply tariff rates to barcodes we require a mapping between product modules and HS codes. We build on the concordance in Bai and Stumpner (2019) that maps NielsenIQ product modules to HS6 codes, forming 359 product categories. The number of product categories is lower than the number of modules due to many cases where several product modules and several HS6 codes have to be combined to form one consistent "category". There are two reasons why this HS6-level concordance is unsuitable for our setting. First, the original concordance matches product module codes to HS6 codes of the 2002 HS revision, while the tariffs relevant to our study were applied using HS codes of the 2017 revision. Second, the concordance links NielsenIQ modules to six-digit HS codes, while tariffs are typically applied at the eight-digit (HS8) or, in some cases, even ten-digit (HS10) level.

To overcome this problem, we translate the concordance in Bai and Stumpner (2019) to the HS2017 revision and refine this concordance, wherever possible, to map to 8-digit or 10-digit HS codes. This results in a more granular concordance between modules and HS codes, creating 445 different categories. Since these categories may map to several HS8 codes, it is possible that the underlying HS8 codes of a given category belong to different tariff waves, creating measurement error via aggregation of heterogeneous tariff levels within a given category. We therefore restrict our attention to a set of “clean” categories, for which all underlying HS8 codes belong to the same tariff wave. Our final sample consists of 200 product categories, of which 39 have been never treated, and 161 belong to wave 3. We then attribute unique tariff rates to the product modules that match to these clean categories, which provides us with a set of 476 product modules, 91 of which have been never treated

and 385 that belong to wave 3. We use these product modules to compute average prices and control for sectoral trends in our analysis further below.[14]

2.5 Sample Construction and Descriptive Statistics

After restricting to a balanced panel of stores and aggregating the sales data to the barcode-by-retailer-by-month level, our sample consists of 787k barcodes in 2017, sold by 68 retailers, and summing up to roughly $219B USD in expenditure in 2017. This is reduced to 206k barcodes in 2017 (out of which 8.4k are Chinese), or 26% of all barcodes in the data once we restrict attention to barcodes with valid firm identifiers and country of origin information. Among others, this stage excludes retailer-owned private label barcodes, since these barcodes are masked by Nielsen and can therefore not be matched to GS1 firm identifiers nor to country of origin information. However, because Label Insight prioritizes products with the highest volume of sales, our matched barcodes account for 63% ($136B USD) of total expenditure in the retail panel.[15] As shown in Figure A.1, the distribution across NielsenIQ departments (both in terms of the number of UPCs and expenditure) looks similar for the matched and broader samples.

3 Tariff Pass-through into Average Prices

We begin by estimating the elasticity of U.S. retail prices with respect to changes in U.S. import tariffs during the 2018-2019 trade war. In order to allow for the broadest possible set of mechanisms associated with endogenous retail price changes, we begin by studying pass-through into aggregated unit values at the firm-module-country-retailer-month level. We denote firms as f, product modules as c, origin countries as j, retail chains as r, and periods as t. Throughout, we refer to the firm-category-country triplet fcj as k to ease notation, such that krt denotes our unit of observation in most estimating equations to follow.

[14] Wave 4a consists of an additional 104 categories that map into 285 product modules.

[15] Flaaen et al. (2025) also study the pass-through of tariffs to consumer prices, focusing on very detailed data on wine imports. Our categories include wine sales, which accounts for $3.9B USD, or 2.8% of all sales in our data.

3.1 Empirical Model

Equation 1 provides our baseline empirical model. In this case y_{krt} denotes some observed outcome associated with the triplet fcj sold at retailer r in month t. We estimate both the effect of tariffs on unit values u_{krt} and quantities sold q_{krt}, where both measures are adjusted for the underlying size of barcodes which constitute each fcj triplet. That is, we construct $q_{krt} = \sum_{i \in \Omega_{krt}} q_{irt}$ where Ω_{krt} denotes the set of barcodes sold by firm f at retailer r in month t within the category-country pair cj. q_{irt} denotes the total quantity sold of barcode i at retailer r in common units such as liters or kgs. We compute this quantity using data on the raw number of packages sold, $\tilde{q}_{irt}$ and the quantity per package, w_i, such that $q_{irt} = \tilde{q}_{irt} w_i$. Moving forward all variables with a tilde denote variables which are unadjusted for product size.[16] Average prices, which we refer to as unit values, are simply expenditure v_{krt} divided by the quantity sold:

$$u_{krt} = \frac{v_{krt}}{q_{krt}} = \frac{\sum\limits_{i \in \Omega_{krt}} p_{irt} q_{irt}}{\sum\limits_{i \in \Omega_{krt}} q_{irt}} = \frac{\sum\limits_{i \in \Omega_{krt}} p_{irt} \tilde{q}_{irt} w_i}{\sum\limits_{i \in \Omega_{krt}} \tilde{q}_{irt} w_i}.$$

We denote the tariff level common to all barcodes within module c and produced in origin country j in period t as τ_{jct}. By applying OLS to Equation 1, we recover the parameter of interest, β:

$$\ln y_{krt} = \beta \ln(1 + \tau_{j(k)c(k)t}) + \alpha_{krt} + \gamma_{k,q(t)} + \epsilon_{krt}. \tag{1}$$

When the dependent variable represents unit values, β captures the elasticity of retail unit values to import tariffs, which we would expect to be bounded below by zero and above by one. An estimate of one would imply that a 10% increase in tariffs generates a 10% increase in retail unit values. When the dependent variable represents total quantity sold, β captures the quantity elasticity of tariffs, which we expect to be negative. For the sake of exposition,

[16] Not all barcodes within a given product module exhibit the same unit of measure. When possible, we convert units into a standard benchmark, such as "OZ" to "mL". Still, some recorded units, such as "Counts", cannot be readily converted to comparable units. For each product module, we calculate the most common unit of measure in our pre-tariff period, 2017, and drop all barcodes with units of measure which cannot be readily converted to this predominant unit of measure. In practice, this data restriction maintains over 95% of expenditure in our raw data. A benefit of using barcodes as our definition of a product is that any change to the size of a barcode, and therefore the labeling, must generate a new barcode such that w_i is constant across all observations of barcode i.

we collect all fixed effects within the vector α_{krt}, such that α_{krt} represents the following set of fixed effects:

$$\alpha_{krt} = \alpha_{kr} + \alpha_{rt} + \alpha_{c(k)t} + \alpha_{j(k)t}.$$

We include fixed effects at the $fcjr$, or kr, level as this is the unit of analysis, such that we control for any level differences in prices and quantities across firm-category-country-retailer units of observation. We then include fixed effects which account for shocks at the retailer-month level, category-month level, and origin country-month level. The category-month shocks account for any demand or supply shocks which are common to barcodes within a given product category. Country-period fixed effects isolate the role specifically of tariffs, by comparing the effect of tariff changes on Chinese goods to other imports from China for which tariffs did not change. Importantly, these fixed effects control for exchange rate changes which may be endogenous to the tariff change as well as any tariff effects which alter perceived future tariff rates associated with all goods produced in China. Finally, we include retailer-period fixed effects to account for potentially differential exposure across retail chains during the trade war.

The vector $\gamma_{k,q(t)}$ constitutes seasonality controls, which we include in order to account for seasonal fluctuations which we observe in the tariff-targeted goods prior to the U.S.-China trade war. These controls constitute a dummy variable for each quarter of the calendar year interacted with the treatment dummy. Such seasonality is not captured by the fixed effects described above as these shocks are not specific to Chinese exports or specific categories, but the specific interaction of Chinese varieties within tariff-targeted categories. We therefore include a vector of seasonality controls, in which $\Gamma_k = 1$ is a treatment indicator equal to one if the combination of product category $c(k)$ and origin country $j(k)$ are ultimately targeted with a tariff by the U.S. We interact this indicator with indicator variables at the quarter level, such that $q(t) = 1$ denotes that period t is in the first quarter.

$$\gamma_{k,q(t)} = \sum_{q'=1}^{4} \gamma_{q'} \times \Gamma_k \times \mathbf{1}[q(t) = q']$$

The identification strategy employed here is effectively a triple-difference: we compare outcomes between Chinese goods and non-Chinese goods across tariff and non-tariff product

categories before and after tariff implementation. As such, we can adjust our specification to an event study so as to check for the presence of pre-trends. To do so, we apply OLS to Equation 2 and plot the estimate vector β_γ.

$$\ln(y_{krt}) = \sum_{\tau=\underline{T}}^{\bar{T}} \beta_\tau \times \Gamma_k \times \mathbf{1}[t = T - \tau] + \alpha_{krt} + \gamma_{k,q(t)} + \epsilon_{krt} \tag{2}$$

T denotes the date of the first tariff change. While our primary focus in this paper are the wave 3 tariffs, we also generate event study figures for product categories targeted by the wave 4a tariffs. In all cases, the control group is the never-treated set of product categories.

In constructing variables at the firm-category-country-retailer level, we restrict ourselves to the set of firm-category-country-retailer quadruplets present in each period of our sample, such that at the krt level we estimate β using a balanced panel. Crucially, as we discuss in Section 4.2, we construct unit values and aggregate quantities using all underlying barcodes within each krt observation, such that while the estimating sample is balanced at the krt level, the underlying set of barcodes used to construct aggregate variables is unbalanced.

3.2 Tariff Elasticity of Retail Unit Values

Table 1 provides our baseline estimates of retail unit value and retail substitution tariff elasticities associated with the wave 3 tariffs applied to imports to the U.S. from China. Columns (1) and (2) represent our preferred estimates, in which the entire sample from January 1 2017 to December 31 2020 is included in our estimation. Column (1) provides the headline result of this paper: retail unit values exhibit a tariff elasticity of 0.435 (0.140). Taken at face value, this estimate implies tariff pass-through into retail prices was 43.5%. To our knowledge, this paper is the first to document a non-zero tariff elasticity of U.S. retail prices for consumer goods during the 2018-2019 U.S.-China trade war. We document significant substitution away from tariff-targeted goods, with a quantity elasticity estimate in Column (2) of -0.610 (0.176). The implied price elasticity of these two estimates is $-(-0.610)/0.435 = -1.40$, which is plausible given the relatively small estimates of price elasticities associated with consumer packaged goods (Faber and Fally 2021) and the relatively short time-horizon we consider here.

Table 1: Tariff Pass-through into Unit Values and Tariff Substitution

	(1) Log Unit Value	(2) Log Quantity	(3) Log Unit Value	(4) Log Quantity
$\log(1+\tau)$	0.435***	-0.610***	0.312***	-0.479***
	(0.140)	(0.176)	(0.097)	(0.130)
Fixed Effects	✓	✓	✓	✓
Exclude 2020			✓	✓
R^2	1.00	0.98	1.00	0.98
Observations	4,291,413	4,291,413	3,220,846	3,220,846

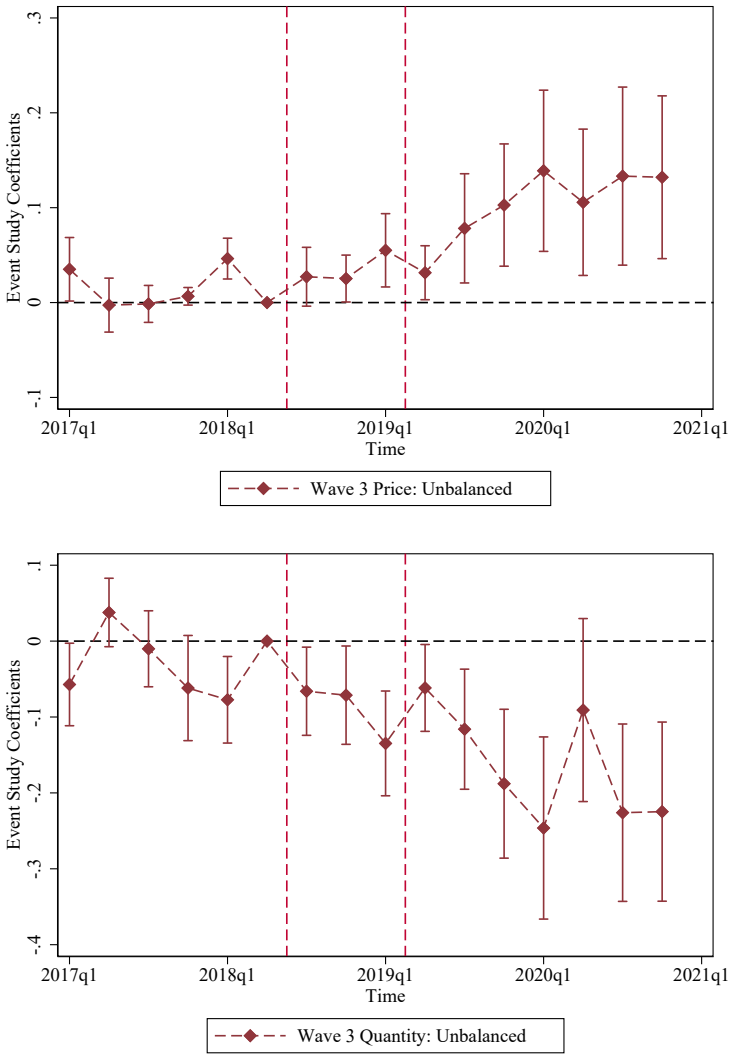
*Notes: This table provides estimates of β from applying OLS to Equation 1. Data are aggregated to the firm-category-country-retailer level. All models include fixed effects by firm-country-category-retailer, retailer-month, category-month, and country-month. Columns (1) and (2) estimate, respectively, pass-through into unit values and tariff substitution in quantities for the entire sample period. Columns (2) and (4) restrict the sample to only time periods prior to 2020, such that our sample is January 1, 2017, to December 31, 2019. Standard errors are clustered three-way at the category, country, and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*

Columns (3) and (4) show that these estimates are broadly robust to excluding the Covid-19 pandemic from our sample. As discussed earlier, the onset of the pandemic in early 2020 could plausibly lead to unobserved demand and supply shocks which may be correlated with treatment. At the same time, the lack of sample periods post-tariffs but prior to the onset of the pandemic limits our ability to recover the dynamic effects of tariffs on unit values and quantities sold when 2020 is removed from our sample. Both the price effects and substitution effects of tariffs are dampened slightly, with the unit values pass-through reduced to 0.312 (0.097) and the quantity elasticity reduced to -0.479 (0.130).

By plotting event study estimates associated with Equation 2, however, we argue that the larger estimates recovered from the full sample are likely not driven by shocks which occurred during the pandemic. Figure 3 provides these event study estimates for unit values (top) and quantities sold (bottom). Each figure includes two vertical lines which denote the two tariff changes associated with tariff wave 3: an initial increase from 2.5% to 12.5% in September 2018 and a further increase to 27.5% in May 2019. There are no appreciable pre-trends in either unit values or quantity units sold prior to the first wave 3 tariff change in September 2018. This is reassuring, as earlier tariff waves (1 and 2) had already occurred, which could have led to anticipatory effects on later tariff waves.

For both unit values and quantities, the initial tariff change exhibits modest effects, with the four quarters after the second tariff increase highlighting the dynamic effect of these

Figure 3: Event Study Estimates for Unit Value Pass-through (Wave 3)



Notes: This figure provides estimates associated with the event study design at the aggregated firm-country-category-retailer level outlined in Equation 2. All models include fixed effects by firm-country-category-retailer, retailer-month, category-month, and country-month. 95% confidence intervals included, with standard errors clustered three-way at the category, country, and month level. Only wave 3 and never-treated categories included. Red vertical lines denote tariff changes, from 2.5% to 12.5% in September 2018, and from 12.5% to 27.5% in May 2019.

tariff changes on price and quantities. Reassuringly, the estimates documented in Figure 3 exhibit a clear stabilization after Q1 of 2020, suggesting the pandemic had little effect on the differential trend of prices and quantities for tariff-affected versus unaffected products. For unit values, this tariff-induced price differential stabilizes at a roughly 12% increase. For quantities, the relative decrease in tariff-affected quantities sold stabilizes at -20%. We

therefore take the wave 3 estimates in Table 1 associated with the entire sample as our headline estimates, such that the tariff elasticity of retail unit values is 0.435.

Robustness across tariff waves. As noted above, the wave 4a tariffs pose a challenge from an identification standpoint, as these tariffs were implemented and then reduced only months prior to the onset of the Covid-19 pandemic. Thus, we focus our analysis in this section on the Wave 3 tariffs, using categories which were never targeted by any wave as control. Still, the results derived here broadly hold for the wave 4a categories as well. Table A.1 confirms that unit values tariff elasticity associated with wave 4a was 0.357 (0.109), such that estimates associated with wave 3 and wave 4a fall within one standard error of each other. Similarly, the quantity elasticity estimate of −0.460 (0.270), while noisier than the corresponding wave 3 estimate, falls within one standard error of the wave 3 estimate. Figure A.2 provides the event study figures associated with wave 4a unit values and quantities, again confirming a lack of pre-trends in both cases.

3.3 Pass-through and the Foreign Cost Share of U.S. Retail

While our unit value tariff elasticity of 0.435 is significant in its own right, as it suggests U.S. consumers bore at least 40% of the incidence associated with these tariffs, our estimate is still well shy of the full pass-through into border prices recorded by numerous papers studying this same episode (Amiti et al. 2019; Fajgelbaum et al. 2020; Cavallo et al. 2021). This is surprising given that our estimate makes use of unit values, thereby offering a plausibly consistent comparison between the unit values recorded in customs data and those recorded in retail prices.[17]

As suggested by Cavallo et al. (2021), any differential in pass-through estimates at the border versus at the store could reflect the ability of U.S. intermediaries, such as retailers, to absorb this cost shock. Given estimates of full pass-through into border prices and null pass-through into store prices, it seems natural to assume that U.S. retailers bore the entire

[17] Flaaen et al. (2020), Cavallo et al. (2021), and Flaaen et al. (2025) do not have quantity sales data for U.S. retailers, and are therefore unable to provide a direct comparison between retail unit values and customs data unit values. Instead, these papers estimate within-product estimates of the tariff elasticity of retail prices, rather than the elasticity of unit values aggregated over multiple products. We discuss this further in Section 4.

incidence of these import tariffs. However this result can only be inferred, as the distribution of incidence across intermediaries and domestic consumers is not directly observed. In this section, we instead consider the possibility that pass-through into retail prices is unlikely to fully reflect the same percentage change in prices as recorded in customs data, even in the presence of full pass-through into border and retail prices. Simply put, tariffs are only levied on the value share of a given imported good when that good crosses the border. All post-tariff costs associated with intra-national distribution and retailing occur after the tariff is levied, such that even full pass-through into border and retail prices would only manifest as the foreign cost share multiplied by the tariff change, when measured using retail prices.[18]

As a simple exercise to highlight this point, consider a retailer importing a good with some unit value u_k^f, as recorded at the border via customs data. This retailer then incurs additional costs associated with post-border services (e.g. transportation, warehousing, etc.) in order to sell this imported good to final consumers. Ignoring second-order substitution effects and assuming a constant pass-through elasticity across the underlying set of barcodes in each kr unit, the effect of a tariff change $d\ln\tau$ on the final price offered at retailer r is simply

$$\beta = \frac{d\ln u_{kr}}{d\ln\tau} = F_{kr}\frac{d\ln u_k^f}{d\ln\tau} = F_{kr}\beta^f.$$

That is, the effect of a tariff τ on the unit values associated with product k as observed at the border, u_k^f, is mediated by the cost share of purchasing this foreign good within the total cost of selling this product on the shelves of retailer r. In this case, F_{kr} denotes this foreign cost share and is defined as $F_{kr} \equiv u_k^f/u_{kr}$. An implication of this first-order cost shock approximation is that estimates of pass-through into retail prices which are lower than estimates of pass-through into border prices may in fact capture the exact same magnitude of tariff pass-through.[19]

Existing estimates of the foreign cost share of U.S. retail imports suggest a range of 30%-50% (Berger et al. 2012). For E.U. wine exports to the U.S., Flaaen et al. (2025) document a

[18] See Sangani (2025) for a similar argument in the context of commodity price shocks and Flaaen et al. (2025) for evidence of this effect in wine tariffs.

[19] If the underlying production function of retailing services is Cobb-Douglas in both the product being purchased and the retailing costs associated with providing this product to consumers, then this result is no long a first-order approximation, but holds exactly for a Cobb-Douglas cost parameter F_{kr} associated with the imported good as input.

foreign cost share of approximately 38%, and provide evidence from the Bureau of Economic Analysis that across consumer goods, the average foreign cost share is just under 50%. We make use of our correspondence between scanner and customs data to estimate the foreign cost share directly at the country-category level in the year 2017, prior to any tariff changes. We therefore impose the restriction $F_{kr} = F_{cj}$, and calculate unit values, or total import value by unit size, at the country-by-HS8 level using U.S. customs data. We then map these unit values into country-by-product module average import unit values, thereby calculating u^f_{cj}: the unit values associated with module c and country j as measured at the border. By calculating the same unit values with retail data, u_{cj}, we construct an estimate of the foreign cost share $\hat{F}_{cj} = u^f_{cj}/u_{cj}$.[20]

We recover an average estimate of $\hat{F}_{cj} = 0.30$, with imports from China exhibiting an average of 0.32. Section B provides details on the construction of the foreign cost share, and Figure B.1 provides a histogram of foreign cost shares that we recover from the data. We manually link our product categories to those studied in Berger et al. (2012) and find that for categories consistent across the two datasets, Berger et al. (2012) recover foreign cost share estimates in the range of 30%-40%, suggesting that our lower estimates may be partly due to the composition of categories we study. Using our estimates of 0.44 (0.31 when restricting attention to pre-Covid data) together with a foreign cost share in the range between 0.3 and 0.5, we infer that pass-through into final prices was likely at least as high as 60% (0.31/0.5).

Still, our estimates of pass-through into retail unit values are considerably larger than the within-product estimates associated with Cavallo et al. (2021), who document zero pass-through for this same tariff wave. Equipped with a unique combination of prices, quantities, and well-defined products, we dedicate the remainder of this paper to understanding this discrepancy. Ultimately, we provide evidence that the composition effect of tariffs on the set of products sold at U.S. retailers is responsible for the majority of tariff pass-through into unit values. We also provide evidence that these compositional effects are welfare relevant, in that they reflect product replacement bias, as in Nakamura and Steinsson (2012), via "shrink-flation": producers and retailers passed through tariffs to U.S. consumers primarily

[20]Further details regarding construction of our measure and descriptions of our findings can be found in Appendix B.1.

by reducing the average package size of barcodes offered. We therefore provide evidence that both the null estimates documented by Cavallo et al. (2021) and the complete pass-through estimates associated with unit values at the border (Amiti et al. 2019; Fajgelbaum et al. 2020; Cavallo et al. 2021) are not necessarily inconsistent, nor does this discrepancy in existing estimates necessarily imply that intermediaries absorbed any substantial share of the cost burden associated with these tariffs.

4 Margins of Adjustment

This section provides a decomposition of the unit value tariff elasticity estimates described above into various margins of adjustment. By incorporating price, quantity, and entry and exit data across barcode-retailer pairs, we decompose unit value pass-through into components associated with the continuing and unbalanced panel of barcodes, and product replacement within the unbalanced panel.

4.1 Decomposing Unit Value Pass-through

In order to fully leverage the detailed price and sales data at our disposal, we begin by providing a decomposition of the tariff elasticity of unit values—total expenditure divided by total quantities sold—as observed at the firm-retailer-module-country-time level. We decompose the aggregate unit value elasticity into two components: unit values associated with continuing barcodes and an adjustment factor that captures the role of entry and exit:

$$u_{krt} = u^c_{krt} \times \frac{u_{krt}}{u^c_{krt}}.$$

We use the superscript c to denote variables associated with this continuing set of barcodes, which we also refer to as the "balanced" panel of barcodes: barcode-retailer combinations observed in every month. The effect of tariffs on the unit value of continuing barcodes is closely related to the effect that is estimated using within-barcode variation for the same set of barcodes. Denoting by $\beta^p_{ir} \equiv \frac{d \ln p_{irt}}{d \ln \tau}$ and $\beta^q_{ir} \equiv \frac{d \ln q_{irt}}{d \ln \tau}$ the (potentially heterogeneous) effects of tariffs on barcode prices and quantities, we can write the effect on the unit value

of continuing goods as follows.[21]

$$\frac{d\ln u^c_{krt}}{d\ln \tau} = \sum_i \left(\frac{v_{irt}}{v^c_{krt}}\right)\left(\beta^p_{ir} + \beta^q_{ir}[1 - u^c_{krt}/p_{irt}]\right).$$

If the effect of tariffs on barcode-level prices and quantities does not differ across barcodes within a firm-retailer-module-country cell, then the effect of tariffs on unit values of continuing barcodes equals the within-barcode effect: $\frac{d\ln u^c_{krt}}{d\ln \tau} = \beta^p$. In general, however, the two elasticities are likely different due to heterogeneous responses across barcodes. For instance, the unit value elasticity would exceed the barcode-level price response if high market share barcodes have a higher price elasticity to tariffs, but would fall short of the barcode-level elasticity if initially more expensive barcodes yielded a larger quantity response. In section 4.4 we estimate price and quantity tariff elasticities within barcode-retailer pairs, therefore providing a comparison with the estimates described in Cavallo et al. (2021), before turning to the role of heterogeneous elasticities across barcodes in section 4.5.

While within-barcode estimates have the benefit of capturing a clear counterfactual price change within the same good by comparing an identical product before and after the tariff change, they are unable to account for the possibility of product replacement bias: the explicit use of product replacement by producers/retailers as a means of raising prices while mitigating the salience of this price change in the eyes of consumers (Nakamura and Steinsson 2012). To the extent that product replacement reflects a genuine change in the price of an otherwise comparable product, this change in price should be seen as relevant for welfare and therefore pass-through.

To capture the role of product replacement, we consider the extent to which unit value tariff elasticities may differ between the balanced and unbalanced panel of barcodes, such that compositional changes in the set of barcodes generate the observed unit value tariff elasticities discussed above. In order to pass on tariffs into higher normalized prices, retailers can either shift the product mix to those products with higher price per unit weight or to products

[21] To derive this equation we write $u^c_{krt} = \sum_i \frac{q_{irt}}{q^c_{krt}} p_{irt}$, so that $\frac{d\ln u^c_{krt}}{d\ln \tau} = \sum_i \frac{v_{irt}}{v^c_{krt}} \left(\frac{d\ln p_{irt}}{d\ln \tau} - \frac{d\ln (q_{irt}/q^c_{krt})}{d\ln \tau}\right)$. Using the fact that $\frac{d\ln q^c_{krt}}{d\ln \tau} = \sum_i \frac{q_{irt}}{q^c_{krt}} \frac{d\ln q_{irt}}{d\ln \tau}$ we get $\frac{d\ln u^c_{krt}}{d\ln \tau} = \sum_i \left(\frac{v_{irt}}{v^c_{krt}}\beta^p_{ir} + \left(\frac{v_{irt}}{v^c_{krt}} - \frac{q_{irt}}{q^c_{krt}}\right)\beta^q_{ir}\right) = \sum_i \frac{v_{irt}}{v^c_{krt}} \left(\beta^p_{ir} + \beta^q_{ir}[1 - u^c_{krt}/p_{irt}]\right)$

with identical sticker price but smaller packaging. To quantify these two different channels, we further decompose the product replacement channel into a channel that captures changes in average sticker prices, and a channel that captures changes in average package size.

We first write quantity as the product of the number of packages sold times package size: $q_{irt} = \tilde{q}_{irt} w_i$. The raw sticker price is then given by the sales volume divided by the number of packages sold, $\tilde{p}_{irt} = v_{irt}/\tilde{q}_{irt}$. Analogously, we therefore calculate the average sticker price as

$$\tilde{u}_{krt} = \frac{v_{krt}}{\tilde{q}_{krt}} = \sum_i \frac{\tilde{q}_{irt}}{\tilde{q}_{krt}} \tilde{p}_{irt}.$$

and average package size as $w_{krt} = \sum_i \frac{\tilde{q}_{irt}}{\tilde{q}_{krt}} w_i$. Unit values can then be expressed as

$$\begin{aligned} u_{krt} &= u^c_{krt} \times \frac{v_{krt}/\tilde{q}_{krt}}{v^c_{krt}/\tilde{q}^c_{krt}} \times \frac{\tilde{q}_{krt}/q_{krt}}{\tilde{q}^c_{krt}/q^c_{krt}} \\ &= u^c_{krt} \times \frac{\tilde{u}_{krt}}{\tilde{u}^c_{krt}} \times \frac{w^c_{krt}}{w_{krt}}. \end{aligned}$$

When differentiated with respect to some tariff change $d \ln \tau$, we derive the following decomposition:

$$\frac{\partial \ln u_{krt}}{\partial \ln \tau} = \frac{\partial \ln u^c_{krt}}{\partial \ln \tau} + \left(\frac{\partial \ln \tilde{u}_{krt}}{\partial \ln \tau} - \frac{\partial \ln \tilde{u}^c_{krt}}{\partial \ln \tau} \right) + \left(\frac{\partial \ln w^c_{krt}}{\partial \ln \tau} - \frac{\partial \ln w_{krt}}{\partial \ln \tau} \right)$$

We turn now to providing tariff elasticity estimates associated with unit values, "sticker prices", and average package size across the unbalanced and balanced panel of barcodes. We then provide a similar analysis within products, thereby providing a full decomposition of the unit value tariff elasticity into constituent components.

4.2 Tariff Elasticity Heterogeneity by Barcode Tenure

This section provides unit value tariff elasticity estimates associated with the balanced and unbalanced panel of underlying barcodes. The balanced panel is constructed using only those barcode-retailer pairs which exhibit non-zero sales in all months of 2016-2020. When estimating tariff pass-through and quantity elasticities for the aggregate unbalanced panel, we again restrict our observations to only those $fcjr$ quadruplets with non-zero balanced panel

expenditure, such that the observations recorded below are identical for the "unbalanced" and "balanced" panel estimates. The terms "balanced" and "unbalanced" instead refer to the underlying set of barcodes used to construct these aggregated measures.

Table 2 provides these estimates for the wave 3 tariffs, with columns (2) and (4) simply repeating the aggregate unit value elasticity estimate and quantity elasticity estimate documented in Table 1. Columns (1) and (3) provide estimates associated with the balanced panel, which are considerably smaller in magnitude, although similar in direction and significance, as the unbalanced panel estimates. Specifically, we recover a unit value tariff elasticity for the continuing set of barcodes of 0.106 (0.025) with a quantity tariff elasticity of -0.262 (0.098). When compared to the unbalanced panel estimates of, respectively, 0.435 (0.140) and -0.610 (0.176), it is clear that pass-through into continuing barcode unit values do not capture the majority of the associated change in unit values across all barcodes. Indeed, the aggregate unit value elasticity is not only over four times larger than the unit value elasticity for continuing barcodes, these estimates are statistically distinguishable from each other at 99% confidence levels. Figure A.3 provides event study figures and confirms that this discrepancy is not due to differential pre-trends. Similarly, Figure A.4 finds a significant discrepancy between event study estimates for the unbalanced and balanced panel for categories associated with the wave 4a tariffs, with the same pattern emerging: continuing product unit values exhibit considerably smaller tariff elasticity estimates when compared to the aggregate unit values elasticities.

4.3 Product Replacement and "Shrink-flation"

This section further decomposes tariff-induced price changes in unit values by barcode tenure into changes in the "sticker price", or unadjusted price per barcode purchased, and a change in the average size of barcodes within a given firm-category-country-retailer quadruplet. We replace the dependent variable in our triple-difference estimating equation—Equation 1—with the "sticker price" unit value $\tilde{u}_{krt}$ and the average package size w_{krt}. We then compare estimates of β across these two variables for the continuing set of barcodes and the unbalanced set of barcodes. As always, we also estimate this triple-difference identification strategy as an event study design, outlined in Equation 2. Figure 4 provides these event study

Table 2: Tariff Pass-through into Unit Values by Barcode Tenure

	(1)	(2)	(3)	(4)
	Log Unit Value		Log Quantity	
$\log(1+\tau)$	0.106***	0.435***	-0.262**	-0.610***
	(0.025)	(0.140)	(0.098)	(0.176)
Balanced	✓		✓	
Unbalanced		✓		✓
Fixed Effects	✓	✓	✓	✓
R^2	1.00	1.00	0.98	0.98
Observations	4,291,413	4,291,413	4,291,413	4,291,413

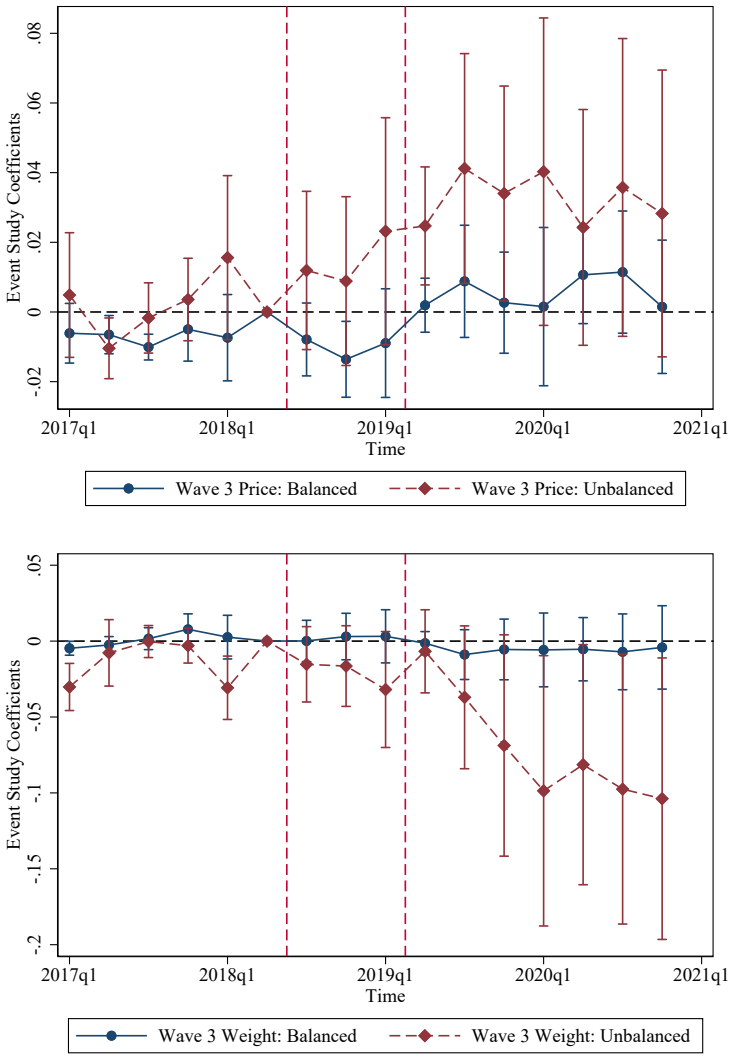
*Notes: This table provides estimates of β from applying OLS to Equation 1. Data are aggregated to the firm-category-country-retailer level. All models include fixed effects by firm-country-category-retailer, retailer-month, category-month, and country-month. Columns (1) and (3) construct unit values and quantities sold using a balanced panel of underlying barcodes with non-zero expenditure in all periods between 2017 and 2020 for each $fcjr$ quadruplet. Columns (2) and (4) construct measures using the raw unbalanced panel of all barcodes recording any sales in a given period. Only $fcjr$ quadruplets with a non-empty set of continuing barcodes are included in estimation, such that both samples have the same observations. Standard errors are clustered three-way at the category, country, and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*

figures, with the blue circle series denoting estimates associated with the continuing set of barcodes, and the red diamond series denoting estimates associated with the unbalanced set of barcodes. The upper panel provides "sticker price" unit values, whereas the lower panel provides average weight estimates.

As shown in the upper panel of Figure 4, average sticker prices increased for the continuing set of barcodes, but only after the additional tariff change to 27.5% in May, 2019, and only by a modest amount. The tariff-induced increase in average sticker prices associated with the unbalanced panel of barcodes exhibits a clear positive response to both tariff changes, with relative prices stabilizing between three and four percentage points higher than the control group after the second tariff change. When placed within the context of the decomposition provided earlier, we therefore have evidence of a modest positive magnitude associated with $\partial \ln \tilde{u}_{krt} / \partial \ln \tau$ and a smaller, albeit positive, magnitude associated with $\partial \ln \tilde{u}^{c}_{krt} / \partial \ln \tau$. Recall from Figure 3, however, that aggregated unit values in the unbalanced panel increased by as much as 12 percentage points after the second tariff change.

The bottom panel of Figure 4 highlights the role of changes in the denominator of unit values, or average weight, in shaping the increase in unit values for the unbalanced panel

Figure 4: Event Study Estimates for Unit Value Pass-through Decomposition



Notes: This figure provides estimates associated with an event study design at the aggregated firm-country-category-retailer level. All models include fixed effects by firm-country-category-retailer, retailer-month, category-month, and country-month. The dependent variable in the upper panel is $\ln \tilde{u}_{krt}$*, expenditure per barcode purchase counts, and in the lower panel the dependent variable is* $\ln w_{krt}$*, the average package size of barcodes purchased. All estimates are associated with wave 3 categories as treated and never-tariff categories as control. Red diamonds denote estimates using the entire sample of underlying barcodes; blue circles denote estimates associated with the continuing set of barcodes. Vertical lines denote tariff changes. 95% confidence intervals included, with standard errors are clustered three-way at the category, country, and month level..*

illustrated in Section 3. As expected, the continuing set of barcodes show no appreciable change in the average size of barcodes sold post-tariff change, as this sample exhibits no entry or exit by construction, and a given barcode's weight is fixed.[22] However the unbalanced

[22]We confirm via the barcode provider GS1 that any change (increase or decrease) to the legally-required declared net content that is printed on the pack, requires assignment of a new GTIN, or barcode, such that

Table 3: Decomposition of Unit Value Pass-through by Barcode Tenure

	(1)	(2)	(3)	(4)	(5)	(6)
	Log Unit Value		Log Sticker Price		Log Average Size	
$\log(1+\tau)$	0.106***	0.435***	0.069**	0.151**	-0.037	-0.284*
	(0.025)	(0.140)	(0.034)	(0.061)	(0.032)	(0.151)
Balanced	✓		✓		✓	
Unbalanced		✓		✓		✓
Fixed Effects	✓	✓	✓	✓	✓	✓
R^2	1.00	1.00	0.99	0.98	1.00	1.00
Observations	4,291,413	4,291,413	4,291,413	4,291,413	4,291,413	4,291,413

*Notes: This table provides estimates of β from applying OLS to Equation 1. Data are aggregated to the firm-category-country-retailer level. All models include fixed effects by firm-country-category-retailer, retailer-month, category-month, and country-month. Columns (1), (3), and (5) construct unit values and quantities sold using a balanced panel of underlying barcodes with non-zero expenditure in all periods between 2017 and 2020 for each $fcjr$ quadruplet. Columns (2), (4) and (6) construct measures using the raw unbalanced panel of all barcodes recording any sales in a given period. Only $fcjr$ quadruplets with a non-empty set of continuing barcodes are included in estimation, such that both samples have the same observations. Columns (1) and (2) estimate pass-through into unit values. Columns (3) and (4) estimate pass-through into "sticker prices", which is total expenditure divided by the un-adjusted number of units sold: $\tilde{u}_{krt} = v_{krt}/\tilde{q}_{krt}$. Columns (5) and (6) estimate tariff elasticities of average package size. Standard errors are clustered three-way at the category, country, and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*

panel exhibits a striking negative trend which begins immediately following the first change in tariffs in September 2018, but intensifies considerably after the second tariff increase in May 2019. By the first quarter of 2020, the average package size of barcodes sold in the unbalanced panel had decreased by ten percentage points when compared to Chinese barcodes in non-tariff categories. In all cases, and in particular for the estimates of average package size, pre-trends are remarkably flat, suggesting that these outcomes are attributable specifically to the import tariffs applied to these goods.

Table 3 then provides the triple-difference estimates associated with these event studies. Columns (1) and (2) simply re-iterate the unit value tariff elasticity estimates associated with, respectively, the balanced and unbalanced panels, as in Table 2. Columns (3) and (4) provide equivalent estimates but for the average sticker price, or price per barcode purchased, with the balanced panel exhibiting a sticker price tariff elasticity of 0.069 (0.034) and the unbalanced panel exhibiting a sticker price elasticity of 0.151 (0.061).

Columns (5) and (6) provide the first clear evidence to date that a remarkable share of

there are no changes in package size within a continuing barcode over time.

tariff pass-through into unit values was in fact achieved by an endogenous decrease in the average package size of goods sold in tariff-affected categories of imports. Specifically, the tariff elasticity of average package size for the continuing set of barcodes is indistinguishable from zero at -0.037 (0.032) while the tariff elasticity of average package size for the unbalanced set of barcodes is -0.284 (0.151), and distinguishable from zero at 90% confidence levels. To put these estimates in context, a 10% tariff increase generates a 2.84% decrease in average package size. This decrease in package size contributed approximately two thirds of the aggregate effect of these tariffs on unit values. The null estimate associated with the continuing set of barcodes suggests that this effect is not attributable to consumer substitution in the face of increasing prices, but rather the endogenous change in varieties offered, due to exit and/or entry.

Replacement bias versus endogenous quality changes. The previous results provide evidence that even within narrow definitions of product—specific firms selling narrow product modules from a given country and at a specific retail chain—product turnover led to a significant increase in the unit values associated with tariff-affected goods, primarily via a decrease in average package size. Whether or not this tariff-induced change in unit values should be considered welfare-relevant, however, remains a first-order concern. A clear benefit to measuring tariff pass-through within continuing products—as in Flaaen et al. (2020), Cavallo et al. (2021), Cavallo et al. (2025), and Flaaen et al. (2025)—is that the researcher directly observes a pre-tariff price associated with each product. The fact that for entering barcodes we have no pre-tariff price is therefore worrisome, as our estimates may conflate endogenous changes in the set of products sold with genuine tariff pass-through via product replacement.[23] Specifically, if tariffs generate an endogenous exit of low-quality, or low-price, products, then the measured unit value may endogenously increase even in the absence of any true tariff pass-through.

Measuring the characteristics of entering and exiting goods over time within the context of Equation 1 is challenging, as entry and exit often do not occur within a given unit

[23]Of course firms may rely on product replacement as a means of increasing prices for the very same reason: it is difficult for consumers to compare the new, more expensive, product directly to the cheaper alternatives which were offered prior to the tariff change and subsequently dropped.

of observation for a given period. We therefore transition to a long-difference estimation strategy in which we take the set of firm-module-country-retailer observations present in Q4 of 2017, prior to any tariff change, and Q4 of 2019, prior to the onset of the covid-19 pandemic, and calculate a number of characteristics regarding entry and exit within each kr observation over this time period. We therefore construct a cross-sectional estimating sample at the kr unit of observation, and regress these measures of entry, exit, and the average package size of entering and exiting barcodes on a simple indicator for whether the kr observation was targeted by wave 3 tariffs. We include category, country, and retailer fixed effects in order to mimic the triple-difference identification strategy from Equation 1. Equation 3 provides our empirical specification with Γ_k denoting an indicator for whether a given firm-country-category pair is treated by a wave 3 tariff, as in Equation 2:

$$\ln y_{kr} = \beta^{LD} \mathbf{1}[\Gamma_k = 1] + \alpha_c + \alpha_j + \alpha_r + \epsilon_{kr}^{LD}. \tag{3}$$

Table 4 provides estimates associated with this regression across a number of different dependent variables. Columns (1) and (2) take the entire sample of continuing kr pairs—those with non-zero expenditure in both Q4 2017 and Q4 2019—and estimate the relationship between wave 3 tariff inclusion and, respectively, exit and entry barcode count shares. If N_{kr0} denotes the initial count of barcodes present in Q4 2017, and N_{kr}^{ex} the count of barcodes which are present in Q4 2017 but not in Q4 2019, we define the exit barcode count share as N_{kr}^{ex}/N_{kr0}. Similarly, the entering barcode count share is N_{kr}^{en}/N_{kr0}, where N_{kr}^{en} denotes the count of barcodes present in Q4 2019 but not Q4 2017.

As the kr level of observation represents a fairly narrow definition of a product, many units do not exhibit any exit or entry over this period. We therefore estimate the effect of tarrifs on these exit and entry shares via PPML in columns (1) and (2). We find a positive effect of tariffs on entry shares and a null effect of tariffs on exit shares. Columns (3) and (4) then narrow the sample to, respectively, only kr pairs which exhibit non-zero exit counts and kr pairs which exhibit non-zero entry counts. We therefore estimate this specification via OLS. The effect of tariffs on entry increases in magnitude while the effect of tariffs on exit becomes positive and significant at levels of 10%, but still smaller in magnitude than

Table 4: Entry and Exit Composition as a Long Difference Cross-Section

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Count Share		Log Count Share		Log Weight		Log Weight	
Wave 3	0.025	0.296**	0.072*	0.413***	0.078**	-0.203***	0.019	-0.305***
	(0.064)	(0.130)	(0.042)	(0.068)	(0.035)	(0.049)	(0.048)	(0.063)
Constant	-1.607***	-1.298***	-1.001***	-0.886***	-0.072***	-0.030***	-0.068***	-0.032***
	(0.004)	(0.009)	(0.003)	(0.004)	(0.002)	(0.003)	(0.003)	(0.004)
Exit	✓		✓		✓		✓	
Enter		✓		✓		✓		✓
Fixed Effects	✓	✓	✓	✓	✓	✓	✓	✓
Cont. Sample	✓	✓	✓	✓	✓	✓	✓	✓
Exit Sample			✓		✓		✓	✓
Enter Sample				✓		✓	✓	✓
Estimator	PPML	PPML	OLS	OLS	OLS	OLS	OLS	OLS
R^2			0.14	0.14	0.07	0.08	0.10	0.09
Observations	162,202	161,747	63,694	54,249	63,694	54,249	35,476	35,476

*Notes: This table provides estimates of the effect of tariffs on product entry and exit characteristics. Data are aggregated to the firm-retailer-category-country level, and estimated as a cross-section, as shown in Equation 3. Columns (1) and (2) estimate the effect of tariffs on the exit count share of barcodes and the entering count share of barcodes using the entire sample. As entry and exit may be zero, PPML is used to estimate columns (1) and (2). Columnns (3) and (4) provide the same estimates but for the sub-sample of kr pairs with non-zero exit and non-zero entry, respectively, via OLS. Columns (5) and (6) estimate the relative average weight of exiting and entering varieties within these non-zero sub-samples via OLS. Columns (7) and (8) provide the same estimates as columns (5) and (6) but for the subset of kr pairs with both non-zero entry and non-zero exit, such that the estimating sample is identical. All estimates contain category, country, and retailer fixed effects. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*

the entry effect.

Columns (5)-(8) estimate the same model but with the average relative weight of exiting and entering barcodes as the dependent variable. For exiting varieties, this dependent variable is $\ln(w_{kr}^{ex}/w_{kr0})$: the sales-weighted average package size of exiting barcodes divided by the average package size in Q4 2017. We construct an identical measure for entering barcodes. Of course, the average weight of entering and exiting barcodes can only be calculated when entry and exit occur, and columns (5) and (6) provide estimates of the effect of tariffs on average package size for the same estimating samples used in columns (3) and (4). Columns (7) and (8) further restrict the sample to only kr pairs with both non-zero exit and non-zero entry, such that the comparison of average barcode weights between exiting and entering barcodes are directly comparable.

For both restricted samples, we find a strong negative effect of tariffs on the average package size of entering barcodes. Both estimates in columns (6) and (8) are statistically

significant at 1%, and both represent the effect of tariffs on average package size of entering barcodes within categories, countries, and retailers. Column (5) provides evidence that exiting barcodes may have been larger in tariff-affected units than control units, this effect is considerable smaller than the negative effect of tariffs on entering barcode size, and disappears entirely when we restrict our sample to kr pairs with both non-zero exit and non-zero entry.

We therefore provide evidence that while tariffs may have induced exit of, on average, somewhat larger barcodes than in control units, tariffs exhibit a strikingly positive effect on barcode entry. This tariff-induced entry is particularly concentrated in varieties with a small package size. The evidence outlined here suggests the unit values tariff elasticity documented earlier is welfare-relevant, and represents product replacement via "shrink-flation": the replacing of barcodes with smaller yet otherwise identical and similarly priced alternatives. This effect occurs within narrow firm-category-country-retailer combinations and is not attributable to increased exit of barcodes with large package size, as would be expected if endogenous quality changes were the driving force behind our estimates.

4.4 Barcode-Level Pass-through

We make use of our barcode-level data to estimate tariff pass-through within barcodes, thereby providing a comparison to the existing retail pass-through literature (Flaaen et al. 2020; Cavallo et al. 2021) and providing a means of comparison with the retail unit value elasticities discussed thus far. The unit of observation throughout is a barcode-retailer pair, such that estimates of pass-through are measured within this same unit of observation.

Empirical Model: We again implement a triple-difference identification strategy reminiscent of the aggregated triple-difference strategy in Equation 1, except now at the barcode-retailer level. We compare barcodes produced in China to non-Chinese goods within the same product category and compare this difference across targeted and non-targeted categories before and after tariffs were implemented, all within the same retail chain. Consider some outcome associated with barcode i at retail chain r in month t, y_{irt}. Equation 4 provides our triple-difference estimating equation at this level of disaggregation, with β capturing the

effect of changes in tariffs on outcome y_{irt} at the barcode-retailer level.

$$\ln(y_{irt}) = \beta \ln(1 + \tau_{j(i)c(i)t}) + \alpha_{irt} + \gamma_{i,q(t)} + \epsilon_{irt} \tag{4}$$

α_{irt} denotes a vector of fixed effects, such that:

$$\alpha_{irt} = \alpha_{ir} + \alpha_{rt} + \alpha_{m(i),t} + \alpha_{j(i),t}$$

This specification replicates the triple-difference identification strategy outlined in Section 3, with the vector $\gamma_{i,q(t)}$ again capturing seasonality controls, as defined earlier. We focus on two primary outcomes of interest at the barcode-retailer pair level: price and total quantity of units sold. We also employ an event study design similar to Equation 2, but at the barcode-retailer level of observation.

Results: We first provide estimates of barcode-level pass-through for a balanced panel of barcodes, such that we omit all barcodes which enter or exit the sample during the tariff period and restrict our analysis to the continuing set of barcodes. Table 5 provides these estimates, and is organized identically to Table 1, except estimates are now interpreted as within barcode-retailer pairs. Columns (1) and (2) provide estimates of tariff pass-through and substitution for the entire estimating sample, and columns (3) and (4) provide the same respective estimates but with 2020 excluded. Similar to the analysis in Section 3, we maintain columns (1) and (2) as our preferred estimates.

Our estimate in column (1) of $\beta = 0.136$ (0.034) implies a tariff elasticity of retail prices of 0.136. When 2020 is excluded, this estimate decreases to 0.112 (0.040). Reassuringly, we find no evidence of pre-trends associated with the price pass-through estimates described in Table 5. Figure A.5 and Figure A.6 in Appendix A.2 illustrate these event study plots for prices and quantities and across both tariff waves. Quantity tariff elasticities at the barcode-retailer level are -0.235 (0.081) for our entire sample period amd -0.361 (0.043) when excluding 2020.

Table 6 provides barcode-level estimates by product tenure, similar to Table 2. When compared to the balanced panel of barcodes, the unbalanced panel exhibits price elasticity

Table 5: Estimates of Tariff Pass-through and Substitution at the Barcode Level

	(1) Log Price	(2) Log Quantity	(3) Log Price	(4) Log Quantity
$\log(1+\tau)$	0.136***	-0.235***	0.112***	-0.361***
	(0.034)	(0.081)	(0.040)	(0.043)
Sample			<2020	<2020
Fixed Effects	✓	✓	✓	✓
R^2	1.00	0.97	1.00	0.98
Observations	1.54e+07	1.54e+07	1.16e+07	1.16e+07

*Notes: This table provides estimates of β from applying OLS to Equation 4. Columns (1) and (2) use the entire sample of observations, with column (1) documenting price pass-through and column (2) the tariff substitution elasticity within barcodes. Columns (3) and (4) provide the same estimates, respectively, but with 2020 excluded from the sample. All estimates are associatd with the continuing panel of balanced barcodes. Standard errors are clustered three-way at the category, country, and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*

Table 6: Estimates of Tariff Pass-through and Substitution at the Barcode Level by Tenure

	(1)	(2)	(3)	(4)
	Log Price		Log Quantity	
$\log(1+\tau)$	0.136***	0.001	-0.235***	-0.689***
	(0.034)	(0.045)	(0.081)	(0.184)
Balanced	✓		✓	
Unbalanced		✓		✓
Fixed Effects	✓	✓	✓	✓
R^2	1.00	1.00	1.00	1.00
Observations	15393284	43819632	15393284	43819632

*Notes: This table provides estimates of β from applying OLS to Equation 4. Columns (1) and (2) provide estimates of, within-barcode pass-through for respectively, the balanced and unbalanced panel of barcodes. Columns (3) and (4) provide the same respective estimates but for the the within-barcode quantity tariff elasticity. Standard errors are clustered three-way at the category, country, and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*

estimates that are almost exactly zero, and precisely estimated. Specifically, we recover within-barcode-retail price elasticity estimates for the unbalanced panel of 0.001 (0.045). The within-barcode-retailer quantity tariff elasticity, however, represents the largest magnitude of substitution throughout our results, at -0.689 (0.184). The implied price elasticity of these unbalanced within-barcode-retailer estimates is implausibly large, suggesting that retailers purposefully kept prices low for barcodes that eventually exited the market.[24]

[24]Our null pass-through results in the unbalanced panel draw an interesting parallel with the null estimate found by Cavallo et al. (2021) who, to the best of our knowledge, similarly estimate pass-through within products using an unbalanced panel of barcode-retailer pairs.

This result is reassuring in that it confirms the role of product turnover in shaping the larger unit value elasticity estimate for the unbalanced panel documented in Section 4.2: this discrepancy is not due to the unbalanced panel of barcodes simply exhibiting larger within-barcode price elasticities than the balanced panel. Indeed the opposite is true: the unbalanced panel of barcodes exhibit precisely zero price effects of tariffs within barcode-retailer pairs. Incorporating all estimates of pass-through recorded up to this point, we recover a striking image of how the typical U.S. consumer would interact with tariff-induced price changes. If a consumer had selected a random unbalanced panel of barcodes and recorded within-barcode price changes throughout this tariff episode, they would recover exactly zero tariff-induced price effects. If they instead focused their efforts on a continuing panel of barcodes, they would recover a tariff elasticity of retail prices of 14%. If they recorded the sales-weighted average price per barcode, such that they interpreted sticker prices as the object of interest, they would recover a tariff price elasticity of 15%. Finally, if they recorded the average price per unit weight of barcodes sold they would recover a tariff unit value elasticity of 44%.

To ensure that the relatively low barcode-level estimates recorded here are credible, we consider a number of robustness checks, including accounting for the possibility of endogenous spillover effects to prices of non-targeted barcodes, tariff "smoothing" by retailers across products, and the effect of tariffs on intermediate goods leading to cost shocks for domestically produced goods. Section B.2 in Appendix B provides detailed descriptions and results for these various robustness exercises: we find no evidence of effects which would suggest that low barcode-level estimates described here are the result of downward bias.

4.5 Heterogeneity in Tariff Pass-through

We now move on to study heterogeneity in pass-through across product and retailer characteristics. In particular, we evaluate the extent to which heterogeneous pass-through by product or retailer market share might lead to a gap between estimates of pass-through within barcodes and estimates of pass-through into average unit values, even in the absence of composition effects.

The key motivation for this exercise is the substantial literature both theoretically ground-

ing, and empirically documenting, heterogeneous pass-through according to market structure and product-specific market power (Atkeson and Burstein 2008; Amiti et al. 2016; Alviarez et al. 2023). In general, these papers either confirm or predict that pass-through should be decreasing in market power, as these firms can reduce mark-ups in order to gain a greater share of the market, whereas small firms are constrained to charging a fixed mark-up over marginal costs. Alviarez et al. (2023), however, suggest that nuance is also needed when mapping pass-through across bilateral buyer-seller relationships, such that both buyer and seller power play a role in shaping the pass-through which occurs when the seller faces a cost shock.

In constructing market shares, we consider three separate definitions of buyer/seller power. In all cases, we use expenditure shares in 2017 to define each measure. First, we define barcode market shares as the within-category expenditure share of a given barcode. Second, we define the seller share, or firm share, of a given firm within a given retail chain as the share of revenue accruing to retailer r generated by sales associated with firm f. Finally, we consider the buyer share, or retailer share, of a given retailer within a given firm, which we define as the share of all revenue accruing to a given firm which is generated within a given retail chain.[25]

Table A.2 provides estimates of tariff pass-through and substitution by barcode market share. We estimate pass-through elasticities which are increasing in market share, with each MS_i denoting elasticity estimates associated with barcodes in the i tercile of 2017 market share. From low to high market share, we recover pass-through estimates of 10.4%, 20.8%, and 23.2%, respectively. Interestingly, tariff quantity elasticities exhibit the reverse pattern, with tariff substitution largest for barcodes in the lowest market share tercile, despite these barcodes exhibiting the lowest level of tariff pass-through.

While tariff pass-through into prices depends on the superelasticity of the demand system, the substitution effect on quantities depends largely on the initial level of the demand elasticity. The results therefore suggest that a larger substitution effect for low market share goods is the result of a very high initial demand elasticity of low market share barcodes – so

[25] 87% of all expenditure on Chinese barcodes is concentrated in the lowest quintile of market share, which hosts 30% of all Chinese barcodes. At the high end of the distribution, 1.6% of all Chinese barcodes are in the highest market share quintile, but these barcodes account for 14% of all expenditure on Chinese barcodes.

high that it overcompensates the smaller increase in prices. Taking our results on prices and quantities together, our results are therefore consistent with Atkeson and Burstein (2008) in the ranking of demand elasticities across market shares (low market share goods have higher elasticities), but inconsistent in the ranking of superelasticities, as our price results suggest that the demand elasticity of high market share products is less elastic, leading to higher pass-through.

We then turn to understanding how pass-through varies by firm and retailer bargaining power via the buyer and seller shares described earlier. For example, Alviarez et al. (2023) build a model of two-sided bargaining according to which pass-through should be increasing in the firm share of retailer sales and decreasing in the retailer's share of firm sales. Figure A.7 plots the distribution of these two bilateral variables in the data in 2017. The blue bars highlight that the vast majority of firms are small relative to retailers: In roughly 80% of all firm-retailer relationships, the firm accounts for less than 0.1% of total retailer sales, and in almost all firm-retailer relationships the firm accounts for less than 1% of retailer sales. This illustrates that there is not much variation in the data to estimate the effect of different firm shares in retailer sales. In contrast, many retailers account for a sizable share of firm sales: In roughly half of all firm-retailer relationships, the retailer accounts for more than 1% of firm sales, and in almost 15% of cases for more than 10%.

We then test for this mechanism in Table 7. We construct two indicator variables based on median cutoffs: one identifying cases where a firm accounts for a high share of a retailer's sales, and another capturing cases where a retailer represents a high share of a firm's sales. Column (1) reports estimates at the barcode–retailer level, while column (2) presents results using unit values as the dependent variable.

Across both specifications, the results highlight the importance of bilateral bargaining power. A larger firm share in a retailer's sales—indicating stronger firm bargaining power—is associated with higher tariff pass-through. Conversely, when retailers hold a stronger bargaining position, pass-through is significantly lower.

We document a positive effect of firm seller power on barcode-level pass-through estimates. Specifically, for a barcode sold by a below-median firm at a below-median retailer, pass-through is 16.7%. For an above-median firm at a below-median retailer, pass-through is

Table 7: Bilateral Sales Shares and Tariff Pass-Through

	(1) Log Price	(2) Log Unit Value
$\log(1+\tau)$	0.167***	0.176
	(0.033)	(0.109)
$\log(1+\tau)\times$ High firm share in retailer sales	0.083***	0.443***
	(0.018)	(0.080)
$\log(1+\tau)\times$ High retailer share in firm sales	-0.163***	-0.295***
	(0.004)	(0.086)
Fixed Effects	ir+mt+jt+rt	fcrj+mt+jt+rt
R^2	0.998	0.996
Observations	15,083,496	4,291,413

*Notes: This table provides estimates associated with interactions between the tariff rate and two measures of market power: a firm's expenditure share within a given retail chain, and a retailer's share of expenditure within a given firm. Column (1) estimates the price elasticity at the barcode-retailer level, whereas column (2) estimates the effect on unit values. Standard errors are clustered three-way at the category, country, and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*

25.0%. However we also document a strong dampening effect of buyer power on barcode-level pass-through, with a below-median firm pass-through that is effectively null when sold at an above-median retailer. When both the firm and retailer are above-median, pass-through is only 8.7%. We therefore match the intuition described in Alviarez et al. (2023): monopoly power exercised by relatively large sellers leads to greater pass-through into barcode-level prices whereas monopsony power exercised by relatively large buyers leads to lower pass-through.

5 Conclusion

This paper provides novel evidence regarding the pass-through and substitution effects of U.S. import tariffs applied to Chinese consumer goods exports during the 2018-2019 U.S.-China trade war. While existing evidence suggests a puzzle, in that tariffs were fully passed through into border prices with little effect on retail prices, we show that in fact such a discrepancy in estimates does not necessarily imply zero pass-through to U.S. consumers. We show that product replacement bias—the replacement of products with higher price per unit weight alternatives—was the predominant mechanism whereby these import tariffs were passed through to U.S. consumers, such that tariff price elasticity estimates within a

given product might be zero while estimates for retail unit values are as high as 0.44. We provide evidence that post-tariff margins associated with mark-ups and/or intra-national distribution and retailing costs imply that a retail price elasticity of 0.44 is not inconsistent with full pass-through into border prices, and implies a likely range of pass-through into retail prices of 60%-100%. Thus, we conclude that U.S. consumers bore the majority of tariff incidence associated with these import tariffs.

Finally, we shed light on the heterogeneous effects of tariffs across barcodes, firms, and retailers, according to buyer and seller power within these producer-retailer relationships. We document increasing pass-through for higher market share barcodes and higher pass-through for firms which constitute a relatively large share of expenditure within a given retailer. Retailers which constitute a relatively large share of expenditure within a given firm, however, dampen the level of pass-through. These effects are present both within barcodes and for our unit value tariff elasticities, but the discrepancy between barcode-specific elasticities and unit values elasticities are particularly pronounced for producers with a high seller share within a given retail chain, such that the extent of heterogeneity we document is dependent on the level of price aggregation in retail data.

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A Appendix: Additional Tables and Figures

This section provides additional tables and figures, as described in the main text.

A.1 Additional Tables

Table A.1: Tariff Pass-through and Substitution by Wave

	(1) Log Unit Value	(2) Log Quantity	(3) Log Unit Value	(4) Log Quantity
$\log(1+\tau)$	0.435***	-0.610***	0.357***	-0.460*
	(0.140)	(0.176)	(0.109)	(0.270)
Wave 3	✓		✓	
Wave 4a		✓		✓
Fixed Effects	✓	✓	✓	✓
R^2	1.00	0.98	1.00	0.98
Observations	4,291,413	4,291,413	3,935,998	3,935,998

*Notes: This table provides estimates of β from applying OLS to Equation 1. Data are aggregated to the firm-category-country-retailer level. All models include fixed effects by firm-country-category-retailer, retailer-month, category-month, and country-month. All columns provide results estimated using the unbalanced panel of underlying barcodes. Columns (1) and (2) estimate the effect of Wave 3 tariffs on unit values and quantities, with all Wave 4a categories omitted. Columns (3) and (4) estimate the effect of Wave 4a tariffs on unit values and quantities with all Wave 3 categories omitted. Standard errors are clustered three-way at the category, country, and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively. Return to Section 3.2.*

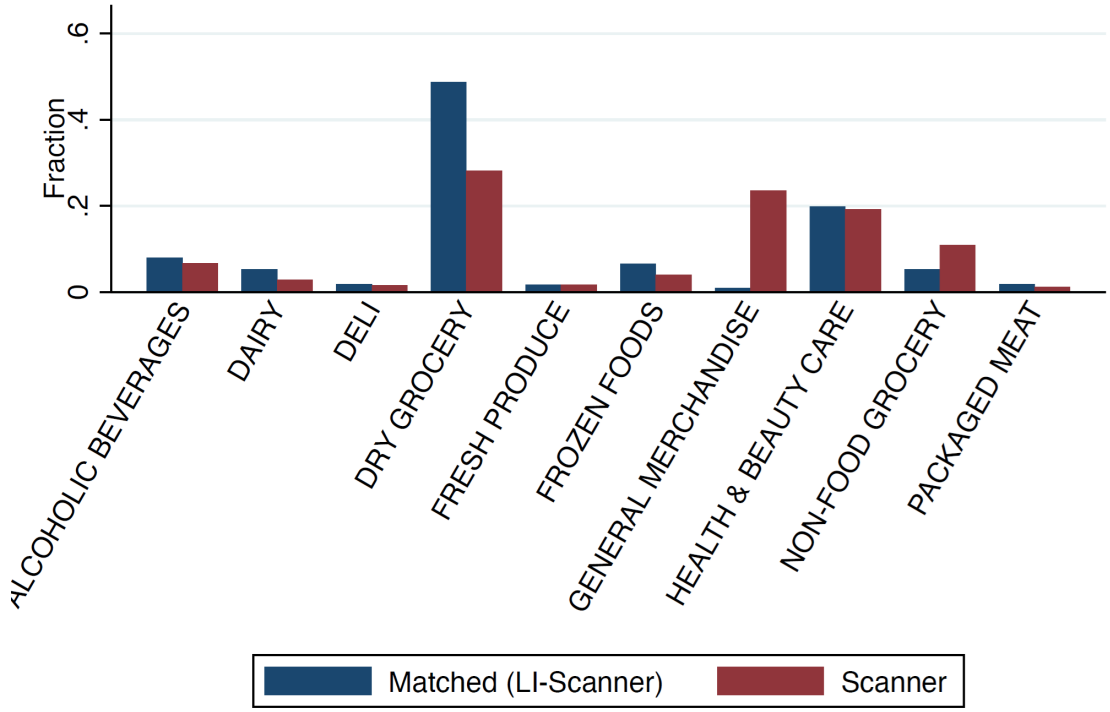
Table A.2: Pass-through and Substitution by Market Share Bins

	(1)	(2)	(3)	(4)
	Log Price		Log Quantity	
$\log(1+\tau) \times MS_1$	0.104**	0.081*	-0.282	-0.397**
	(0.039)	(0.044)	(0.219)	(0.191)
$\log(1+\tau) \times MS_2$	0.208***	0.174***	-0.090	-0.280
	(0.050)	(0.054)	(0.214)	(0.196)
$\log(1+\tau) \times MS_3$	0.232***	0.217***	-0.197	-0.253
	(0.065)	(0.058)	(0.310)	(0.256)
Sample		<2020		<2020
Fixed Effects	✓	✓	✓	✓
R^2	1.00	1.00	0.97	0.98
Observations	15393284	11556925	15393284	11556925

*Notes: Tariff pass-through by barcode market share, from low market share (s_1) to high market share (s_3). Market shares are computed within product modules based on 2017 sales volume. Columns (1) and (3) estimate price and quantity elasticities for the entire sample, with columns (2) and (4) providing the same respective estimates with 2020 excluded. Standard errors are clustered three-way at the category, country, and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively. Back to Section 4.5.*

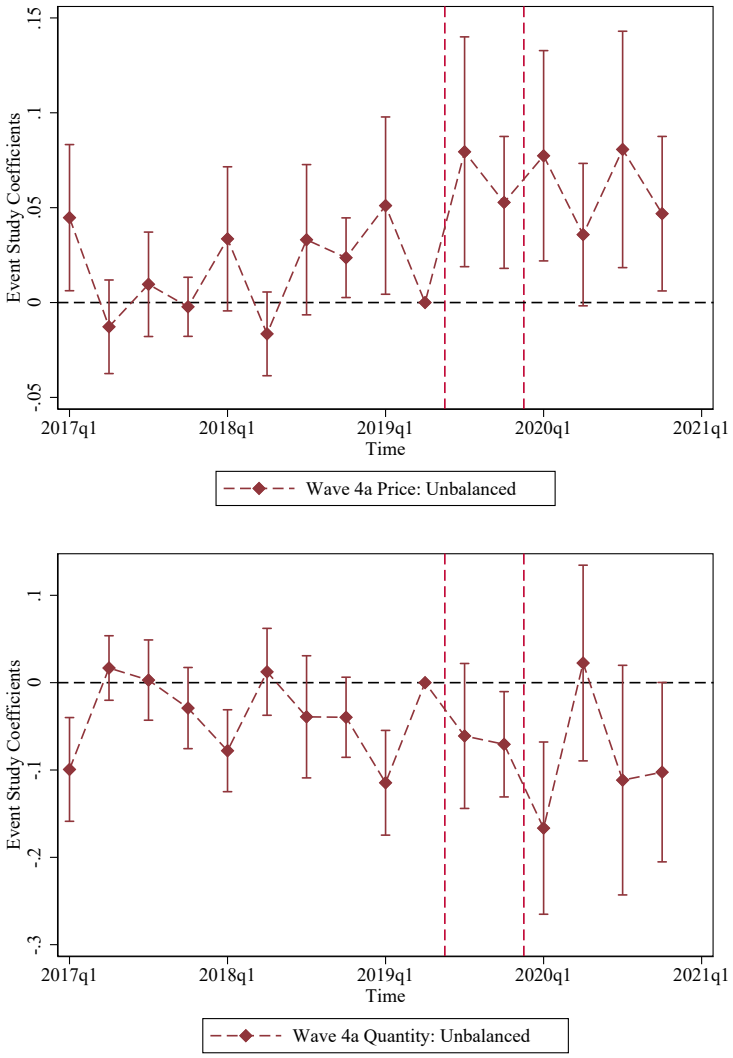
A.2 Additional Figures

Figure A.1: Distribution of Barcodes by Department



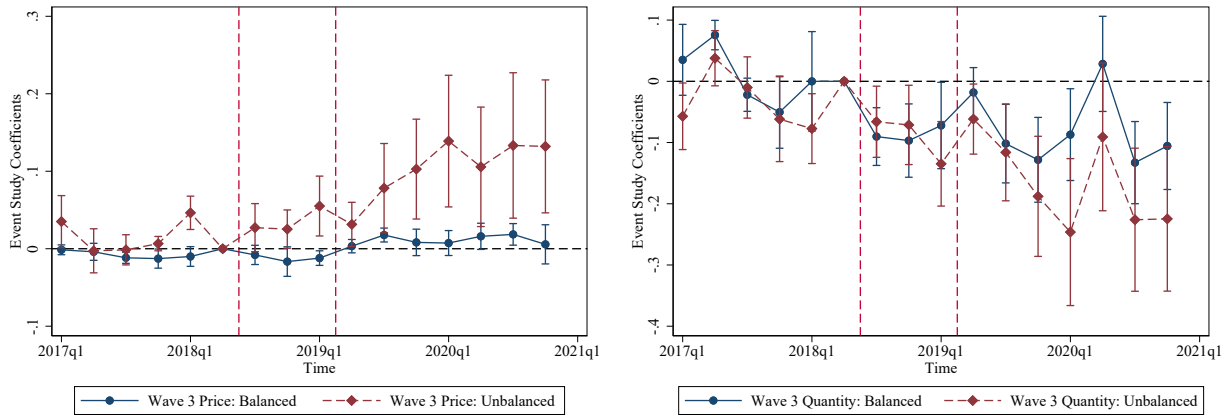
Notes: Figure 1 provides the distribution of barcodes across NielsenIQ product departments according to the raw scanner data (red) and the merged Label Insight-NielsenIQ dataset used in this paper (blue).

Figure A.2: Event Study Estimates for Unit Value Pass-through (Wave 4a)



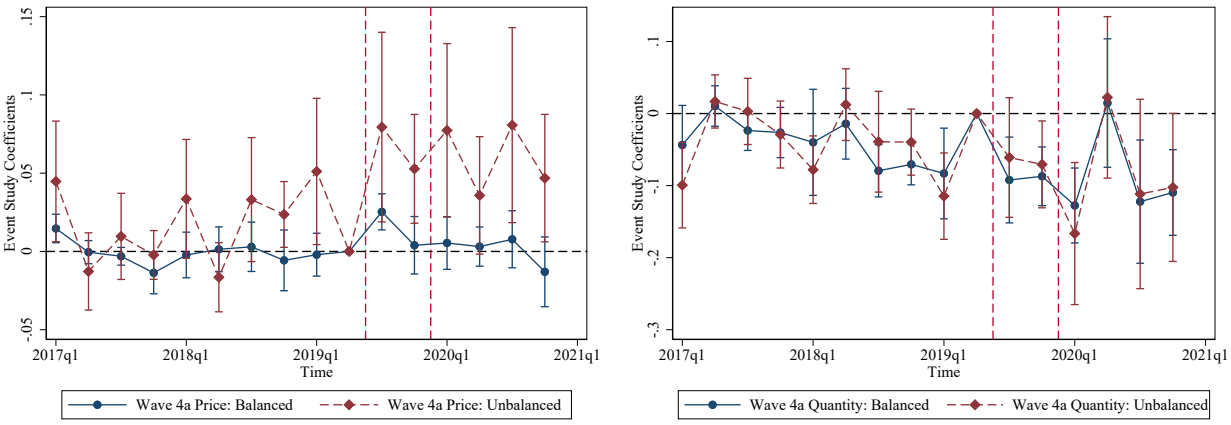
Notes: This figure provides estimates associated with the event study design at the aggregated country-category-retailer level outlined in Equation 2. All models include fixed effects by firm-country-category-retailer, retailer-month, category-month, and country-month. 95% confidence intervals included, with standard errors clustered three-way at the category, country, and month level. Only wave 4a and never-treated categories included. Red vertical lines denote tariff changes, from 5.0% to 20.0% in August 2019, and from 20.0% to 12.5% in January 2020. Back to Section 3.2.

Figure A.3: Event Study Estimates for Unit Value Pass-through by Barcode Tenure



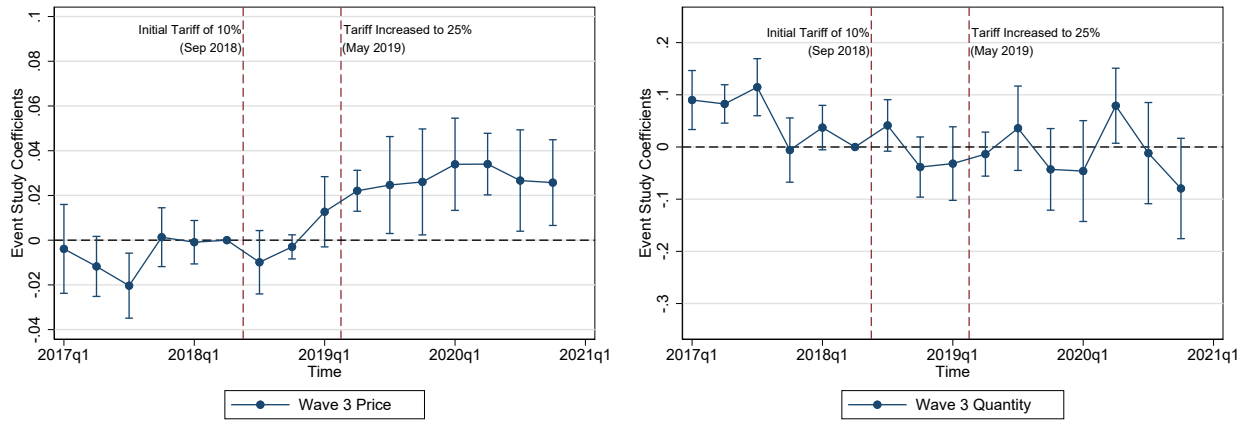
Notes: This figure provides estimates associated with the event study design at the aggregated country-category-retailer level outlined in Equation 2. All models include fixed effects by country-category-retailer, retailer-month, category-month, and country-month. 95% confidence intervals included, with Standard errors clustered three-way at the category, country, and month level.

Figure A.4: Event Study Estimates for Unit Value Pass-through by Barcode Tenure (Wave 4a)



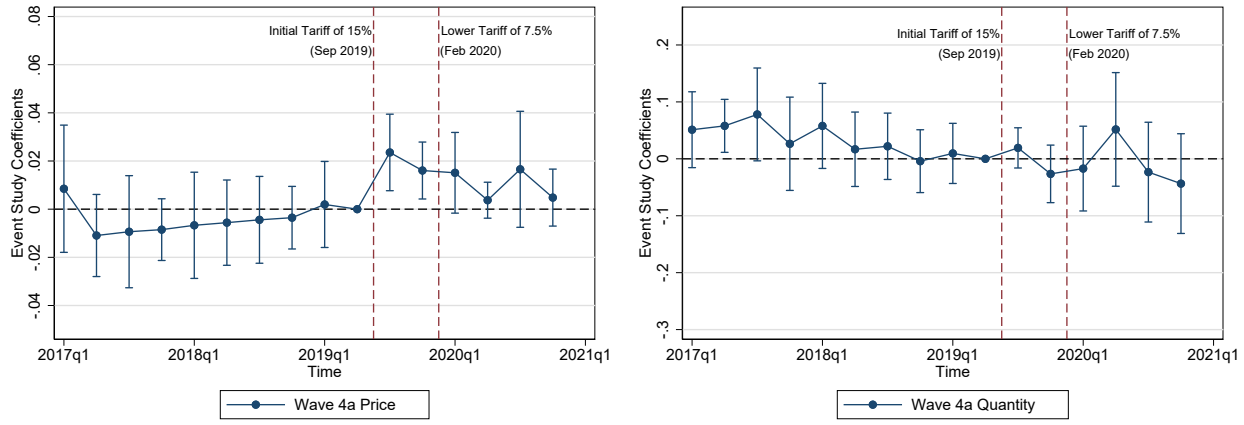
Notes: This figure provides estimates associated with the event study design at the aggregated firm-country-category-retailer level outlined in Equation 2. All models include fixed effects by firm-country-category-retailer, retailer-month, category-month, and country-month. 95% confidence intervals included, with Standard errors clustered three-way at the category, country, and month level.

Figure A.5: Event Study Estimates at Barcode Level (wave 3)



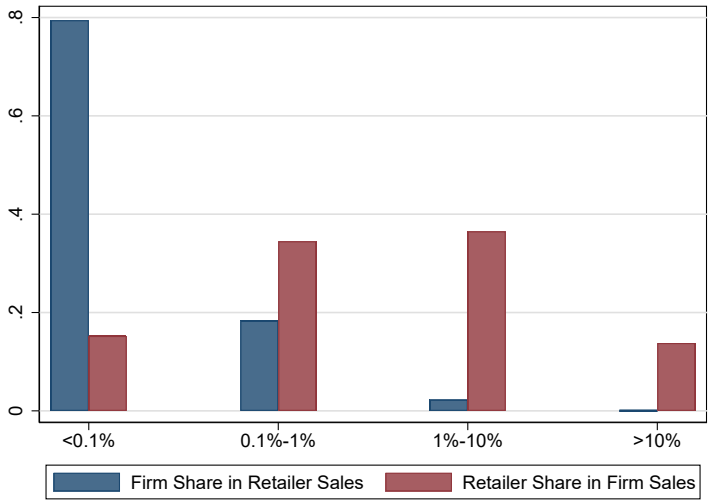
Notes: This figure provides estimates associated with a barcode-retailer-level event study. The dependent variable is p_{irt} (left) and q_{irt} (right). All never-tariff categories are included as the control group, whereas the only treated group included are those categories targeted by tariff wave 3. Barcode-retailer, retailer-month, category-month, and country-month fixed effects included. 95% confidence intervals included, with Standard errors clustered three-way at the category, country, and month level.

Figure A.6: Event Study Estimates at Barcode Level (wave 4a)



Notes: This figure provides estimates associated with estimating Equation 4. The dependent variable is p_{irt} (left) and q_{irt} (right). All never-tariff categories are included as the control group, whereas the only treated group included are those categories targeted by tariff wave 4a. Barcode-retailer, retailer-month, category-month, and country-month fixed effects included. 95% confidence intervals included, with Standard errors clustered three-way at the category, country, and month level.

Figure A.7: Firm and Retailer Bilateral Dependence



Notes: Distribution of the firm share in retailer sales and the retailer share in firm sales in 2017. Both variables are computed for 2017 data. Firm share in retailer sales is the total sales by products of firm f in retailer r, divided by total sales by retailer r. Retailer share in firm sales is the total sales by products of firm f in retailer r, divided by total sales by firm f.

B Empirical Appendix

This section provides further detail regarding data construction, empirical methodology, and results.

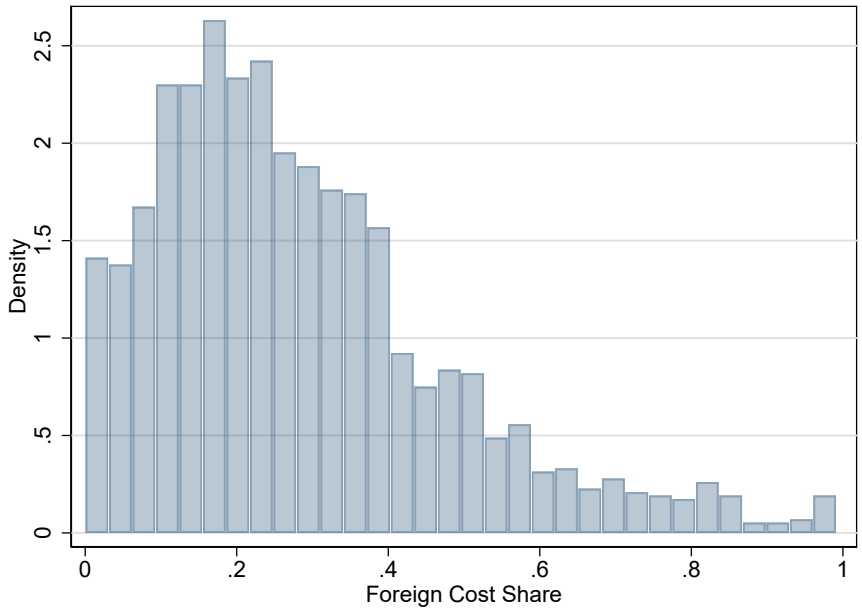
B.1 Constructing Foreign Cost Share Estimates

While we do not have data on the border prices of individual barcodes, we make progress by comparing category-by-origin unit values between retail and customs data. Since the U.S. computes tariff duties using free-on-board (FOB) prices, we use the HS10 customs valuation unit values by country of origin, as opposed to cost-in-freight (CIF) unit values which include international transport and insurance. We drop observations with invalid or missing quantity units and aggregate the import data to the level of product category-by-origin-by-unit of measurement, and compute unit values at that level.

We then try to harmonize as much as possible quantity units of measurement in both the scanner and the customs data. Some unit values cannot be compared across the two datasets, for instance if the unit in the scanner data is simply a "count", but it is measured in kilograms in the customs data. For the pairs for which we can identify a common quantity unit, we compute the foreign cost share as the ratio of the customs unit value and the scanner unit value. In rare cases, there are multiple quantity units of measurement in a given category-by-origin country cell. In all of these cases, one of the units is a product count while the other one is a more cleanly defined unit, such as volume or weight. In these cases we prefer the unit value measured from the non-count data. This procedure is arguably imperfect, since we cannot be sure that we compare the same barcodes in the customs and scanner data. Still, within the roughly two thousand country-category pairs we merge between datasets, 95% of all recovered estimates of the foreign cost share are less than one, which is reassuring as the theoretical upper bound of the FCS is one. Reassuringly, we recover an unweighted correlation between unit values in scanner data and customs data of 0.70. Figure B.1 provides a histogram of the foreign cost share estimates we recover across all imported country-category pairs.

We find no clear relationship with the foreign cost share estimates we recover and import

Figure B.1: Distribution of Foreign Cost Share Estimates



Notes: This figure provides a histogram of our foreign cost share estimates for all estimates which we reliably link between customs data and scanner data. $N = 1852$.

origin characteristics, such as distance to to the U.S. or GDP per capita, conditional on the product category in question. We do find some evidence that foreign cost shares associated with specifically with imports from China may have within-category foreign cost shares which are, on average, 3.6 percentage points lower than average. Still, this evidence remains qualitative at best, given the limited number of observations.

B.2 Robustness of Barcode-Level Estimates

This section describes a number of robustness checks associated with spillover effects of tariffs to non-treated barcodes. First, any model of strategic pricing would suggest that tariffs on Chinese varieties might induce an increase in markups for competing varieties from other origin countries. We test this effect directly using a differences-in-differences identification strategy, and comparing the effect of tariffs on the price of non-targeted products within tariff categories to products in non-tariff categories. Specifically, we estimate the following

Table B.1: Mark-ups of Non-Tariff Goods

	(1)	(2)	(3)	(4)
	log Price			
$\log(1+\tau_{CN})$	0.001	-0.001	0.019	0.011
	(0.021)	(0.019)	(0.026)	(0.023)
$\log(1+\tau_{CN}) \times (Share_{CN} > 0)$			-0.032	-0.020
			(0.025)	(0.022)
Sample		<2020		<2020
Fixed Effects	✓	✓	✓	✓
R^2	1.00	1.00	1.00	1.00
Observations	1.51e+07	1.13e+07	1.51e+07	1.13e+07

*Notes: This table provides estimates of β from applying OLS to Equation B.1. In columns (1) and (3) we estimate our specification on the whole sample, while in columns (2) and (4) we restrict attention to the comparison between wave 3 and the never treated, and stop the sample pre-Covid. Standard errors are clustered two-way at the category-country and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*

empirical model:

$$\ln(p_{irt}) = \beta \ln(1+\tau^{cn}_{c(i)t}) + \alpha_{ir} + \alpha_{rt} + \alpha_{j(i)t} + \alpha_{g(i)t} + \gamma_{w(i)q(t)} + \epsilon_{irt} \tag{B.1}$$

where $\tau^{cn}_{c(i)t}$ denotes the tariff applied to Chinese varieties within category c, which is now the explanatory variable for all non-Chinese varieties within that same category. All other fixed effects remain the same, other than the module-time fixed effects. Since we now exclude Chinese products from the estimation, these fixed effects would entirely absorb the tariff variation across categories. To still control for differential sector-level price trends we include fixed effects at a higher level of category classification, $\alpha_{g(c)t}$, where $g(c)$ denotes roughly 100 NielsenIQ product groups. Results are in Table B.1 and do not provide evidence for an endogenous mark-up response of non-tariff-affected goods. Estimated coefficients are all close to zero with relatively tight confidence intervals. It should be noted that in a similar exercise, Cavallo et al. (2021) also find little evidence for anti-competitive behavior within consumer goods. However, within consumer packaged goods, China's expenditure share in the U.S. is smaller than in many other export markets, with shares that are typically less than 5% of all sales. Thus, it is not immediately clear that the combination of low market shares and (relatively) small tariff changes would lead to observable anti-competitive behavior.

Second, we consider the possibility that firms or retailers may choose to spread price increases across many goods instead of concentrating them on the taxed items. While there is some anecdotal evidence for this type of strategy, particularly by retailers, we are not aware of any study testing whether this is a common pattern.[26] We test this mechanism as follows: do retailers (or firms) more exposed to a particular tariff wave tend to increase their prices of non-taxed goods by more than other retailers? To avoid retailers (or firms) being treated multiple times through different waves, we again restrict attention to the treatment through wave 3, and stop the sample in December 2019. We then define a set of treated firms and retailers using information on the firm and retailer product portfolio in 2017. Among the roughly 4,200 firms in our sample, we define a firm as treated if in 2017 it sold some Chinese products that were later taxed through Wave 3. This applies to 188 firms. Since all 67 retailers have some level of exposure to Chinese Wave 3 goods, we divide retailers by the median 2017 exposure. We then estimate the following specification on only non-Chinese barcodes,

$$\ln(p_{irt}) = \beta \text{Treated Retailer}_r \times \text{Post}_t + \alpha_{irt} + \epsilon_{irt}, \tag{B.2}$$

where $textPost_t$ is a dummy that equals one after the implementation of the initial Wave 3 tariffs in September 2018.

Since we want to identify the effect using variation across retailers, we now need to suppress the retailer-by-time FE. In column (1) of Table B.2 we estimate equation B.2 using barcode-by-retailer, country-by-time, and module-by-time effects, and we do the same in column (2) when we look at the firm effect. Column (3) adds both effects in the same specification. Column (4) then implements a more demanding specification to estimate the firm effect, where we suppress all cross-retailer variation. Finally, specification (5) implements a more demanding specification to estimate the smoothing effect for retailers, using only variation for the same barcode sold at the same point in time, across different retailers, which we do by adding a barcode-by-time effect.[27]

[26] As an example: https://www.oliverwyman.com/our-expertise/insights/2025/jan/retail-strategies-to-adapt-to-new-us-tariffs.html

[27] Notice that a barcode-by-time effect wipes out the firm smoothing effect, since a barcode always remains in the same firm.

Table B.2: Smoothing of Price Increases

	(1)	(2)	(3)	(4)
	Log Price			
Treated Retailer × Post	0.004**		0.004**	
	(0.002)		(0.002)	
Treated Firm × Post		0.003	0.003	0.002
		(0.004)	(0.004)	(0.004)
Sample	<2020	<2020	<2020	<2020
Fixed Effects	ir,mt,jt	ir,mt,jt	ir,mt,jt	ir,mt,rt,jt
R^2	1.00	1.00	1.00	1.00
Observations	1.51e+07	1.51e+07	1.51e+07	1.51e+07

*Notes: In column (1) we estimate equation B.2 using barcode-by-retailer, country-by-time, and module-by-time effects, and we do the same in column (2) when we look at the firm effect. Column (3) adds both effects in the same specification. In column (4) we estimate a more demanding specificaiton with retailer-by-time FEs. Standard errors are clustered two-way at the category-country and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*

Table B.2 provides some evidence that, among non-taxed goods only, retailers that were initially more exposed to tariffs raised their prices by more, suggesting that retailers did smooth tariff increases. However, the magnitude is quite small, with prices of unaffected goods increasing by only 0.8-0.9% after implementation of Wave 3. While this channel is of interest *per se*, it is certainly not strong enough to explain the relatively low barcode-level pass-through estimates we recover of 10%.

Finally, while many consumer goods were taxed in waves 3 and 4 of the 2018/19 trade war, earlier waves (1 and 2) focused on intermediate goods. This might lead to price increases of domestically produced consumer goods to the extent that those goods use taxed imported intermediates as inputs in production. Since we cannot directly observe imported intermediates at the firm-level in our data, we instead use the fact that this concern only applies to U.S.-produced goods in the control group, but not to foreign-produced goods. In Table B.3, we re-estimate Equation 4, excluding U.S.-produced goods entirely. The coefficients stay very close in magnitude to our main estimates, suggesting that imported intermediates are not a first-order concern and do not bring us closer to estimates of pass-through in the range of 40%, which we recovered in Section 3.

Table B.3: Comparison to other Imported Products

	(1)	(2)
	Log Price	
$\log(1+\tau)$	0.101**	0.093**
	(0.044)	(0.042)
Sample		<2020
Fixed Effects	✓	✓
R^2	1.00	1.00
Observations	2.76e+06	2.07e+06

*Notes: This table provides estimates of β when applying OLS to equation 4, excluding US-made goods from the estimation. Standard errors are clustered two-way at the category-country and month level. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.*