Exchange Rates, Product Variety, and Substitution in U.S. Scanner Data*

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Working Paper
January 30, 2019

Abstract

This paper is the first to combine U.S. retail scanner data with barcode-level data on country of origin. We use this novel dataset for 10 narrow product categories from 2001 to 2012 to study several related questions. First, we estimate exchange rate pass-through into the relative CES price index of foreign retail products to be about -0.25 for the median product category. Consistent with the strong positive correlation between the number of foreign products and the dollar that we document, we estimate that exchange rate pass-through is about half the previous magnitude when entry/exit effects are excluded from the price index. Further, our results using a structural demand system provide clear evidence that the macro elasticity of substitution (between U.S. and foreign goods) is significantly smaller than the micro elasticity (within U.S. and foreign goods) for essentially all product categories. Finally, we find little evidence of income-induced expenditure switching.

JEL CLASSIFICATION: D12, E31, F31
KEYWORDS: exchange rate pass-through, product variety, elasticity of substitution, expenditure switching, non-homotheticity

*Any opinions and conclusions expressed herein are those of the authors and do not necessarily represent the views of the Federal Reserve Board or anyone affiliated with the Federal Reserve System. All estimates and analyses based on Information Resources Inc. data are by the authors and not by Information Resources Incorporated. We also gratefully acknowledge using data from Label Insight.

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1 Introduction

The responsiveness of import prices to exchange rate changes is a key mechanism through which monetary policy affects the macroeconomy. Further, the real effect of such exchange rate changes is determined by the extent to which consumers switch expenditure between domestic and foreign products in response to relative price changes. For example, as the dollar appreciates against foreign currencies, imported goods become relatively cheaper for U.S. consumption. The extent to which consumers will shift their consumption towards these relatively cheaper, imported products depends on the substitutability of domestic and foreign products.

A large previous literature has shown that the pass-through of exchange rates into prices varies by country and product, and is typically incomplete, meaning that the change in prices is less than the change in exchange rates. For example, Campa and Goldberg (2005) estimate that the United States has one of the lowest measures of exchange rate pass-through into import prices in the OECD, at about -0.25 in the short run and -0.40 in the long run[1]. In contrast, Nakamura and Steinsson (2012) argue that product entry and exit creates a bias in the U.S. import price index, and therefore exchange rate pass-through is actually higher than standard estimates suggest[2]. Going beyond import prices, recent disaggregated analysis has estimated that the long run exchange rate pass-through into U.S. beer retail prices is approximately -0.05 to -0.10, with limited pass-through accounted for by retailers' markup adjustments, local non-traded costs, and brand price adjustment costs (Goldberg and Hellerstein (2012)).

These price responses to exchange rates can induce expenditure-switching between domestic and foreign products, depending on their substitutability. Feenstra et al. (2018) compare the elasticities of substitution within foreign-origin products (micro elasticity) to the elasticities of substitution between U.S. and foreign products (macro elasticity) for broad product categories, such as food and apparel manufacturing. They estimate the micro elasticity to be larger than the macro elasticity in about half of their product categories, but cannot reject that the micro and macro elasticities are equal for the other half of their product categories. On the other hand, recent research using disaggregated scanner data in Latvia suggests that expenditure-switching can also be driven by income effects, rather than relative prices (Bems and Di Giovanni (2016)).

In this paper, we combine U.S. retail scanner data with barcode-level country of origin data, allowing us to identify the foreign-origin barcodes on U.S. store shelves for the first time[3]. We use this novel dataset for 10 narrow product categories from 2001 to 2012 to estimate a nested, non-homothetic demand system. Our results provide uniquely clean and novel evidence that support the views of Nakamura and Steinsson (2012) and Feenstra et al. (2018), while challenging the applicability of the results of Bems and Di Giovanni (2016) to a country like the United States.

We first use an identification approach based on Feenstra (1994) to estimate parameters for the (micro) elasticities of substitution within domestic and foreign products separately in

[1] An exchange rate pass-through of -0.25 means that a 10 percent appreciation of the dollar results in a 2.5 percent decrease in prices.

[2] However, see Gagnon et al. (2014) for an opposing view.

[3] This has not been previously possible in U.S. data, with the exception of beer as in Hellerstein (2008) and Goldberg and Hellerstein (2012). Similar data have been used for the UAE in Antoniades and Zaniboni (2016), and for Switzerland in Auer et al. (2018).

each product category. For about half of our product categories, the within elasticity for foreign products is smaller than the elasticity for domestic products. For the other half of our categories, the within elasticity parameters are about the same for foreign and domestic products. Overall, our median estimate of the within elasticity for domestic products is about 5.2, while our median estimate for foreign products is about 4.4. Using these estimates and the barcode-level price data, we then construct constant elasticity of substitution (CES) price indexes for foreign and domestic products in each product category and store in the scanner data.

Second, we use store-category, share-weighted exchange rate indexes to estimate exchange rate pass-through into the CES price index of foreign relative to domestic products. For the median product category, we estimate exchange rate pass-through of about -0.25. We document a strong correlation between the number of foreign products on store shelves and the exchange rate, such that dollar appreciation results in more foreign products on store shelves. When we re-estimate exchange rate pass-through excluding the effect of product entry and exit from the CES price index, we find that pass-through is about half its previous magnitude.

Third, we use the changes in the relative price of foreign to domestic products that result from exchange rate changes to estimate the (macro) elasticity of substitution between foreign and domestic products in each product category. We find that the between elasticity is significantly smaller than the within elasticity for essentially all product categories. In particular, our median estimate of the between elasticity is about 2.7. However, we find little evidence of income-induced expenditure switching.

The rest of this paper is structured as follows. In Section 2, we develop our nested demand system. Section 3 outlines our identification strategy for estimating the within and between elasticities of substitution. In Section 4, we describe our data in detail. We report and analyze our estimation results in Section 5. Section 6 concludes.

2 Model

We will be modeling the consumer choice among foreign and domestic versions of products that are available in the same store. In our structural model, product categories (such as beer and cereal) are represented by groups g, stores are represented by s, and barcodes or universal product codes are represented by u. We have products of foreign origin f and domestic origin d. In our notation, a variety refers to a barcode-store pair such that $u \epsilon B_{dsgt}$ represents a barcode u in the set of all domestic barcodes sold in group g and store s at time t.

For every product category, we have a nested preference structure such that the upper-tier utility in a store-category is given by the non-homothetic generalization of CES preferences of Hanoch (1975), as formulated in Comin et al. (2015). This upper-tier utility is defined by

$$\varphi_{dsgt}^{\frac{1}{\sigma_g^b}} Q_{dsgt}^{\frac{\sigma_g^b-1}{\sigma_g^b}} U_{sgt}^{\frac{\epsilon_g^d-1}{\sigma_g^b}} + \varphi_{fsgt}^{\frac{1}{\sigma_g^b}} Q_{fsgt}^{\frac{\sigma_g^b-1}{\sigma_g^b}} U_{sgt}^{\frac{\epsilon_g^f-1}{\sigma_g^b}} = 1 \tag{1}$$

where utility U in store-category sg is implicitly defined, σ_g^b is the (macro) elasticity of substitution between domestic and foreign barcodes of product category g, ϵ_g^d is a measure of the income elasticity of domestic products in category g, and φ_{dsgt} is a taste shifter for domestic goods in store-category sg at time t. Similarly, ϵ_g^f is a measure of the income elasticity of foreign products in category g, and φ_{fsgt} is a taste shifter for foreign goods in store-category sg at time t.

The lower tier of the preference structure is such that the domestic and foreign products in category g at time t are CES aggregates given by

$$\begin{aligned} Q_{dsgt} &= \left[\sum_{u \epsilon B_{dsgt}} (\varphi_{ut} Q_{ut})^{\frac{\sigma_d^w}{\sigma_d^w - 1}} \right]^{\frac{\sigma_d^w}{\sigma_d^w - 1}} \\ Q_{fsgt} &= \left[\sum_{u \epsilon B_{fsgt}} (\varphi_{ut} Q_{ut})^{\frac{\sigma_f^w}{\sigma_f^w - 1}} \right]^{\frac{\sigma_f^w}{\sigma_f^w - 1}} \end{aligned} \tag{2}$$

where σ_d^w is the (micro) elasticity of substitution within domestic barcodes of product category g, σ_f^w is the (micro) elasticity of substitution within foreign barcodes of product category g, and φ_{ut} is a taste shifter for barcode u in store-category sg at time t. Our nested preference structure is therefore a generalization of the Nested CES preferences used in Feenstra et al. (2018).

Solving the consumer problem by backward induction, the utility-maximizing quantity demanded of a domestic variety $u \epsilon B_{dsgt}$ is given by

$$Q_{ut} = \left[\frac{P_{ut}^{-\sigma_d^w} \varphi_{ut}^{1-\sigma_d^w}}{\sum_{j \epsilon B_{dsgt}} (\frac{P_{jt}}{\varphi_{jt}})^{1-\sigma_d^w}} \right] Y_{dsgt} \tag{3}$$

where Y_{dsgt} is given expenditure on domestic store-category sg at time t and P_{ut} is the price of variety u at time t. Similarly, we have an analogous equation for the quantity demanded of a given foreign variety.

The price index for domestic products in store-category sg at time t is given by the following CES price index

$$P_{dsgt} = \left[\sum_{j \epsilon B_{dsgt}} \left(\frac{P_{jt}}{\varphi_{jt}} \right)^{1-\sigma_d^w} \right]^{\frac{1}{1-\sigma_d^w}} \tag{4}$$

where the change from time $t-i$ to time t is

$$\frac{P_{dsgt}}{P_{dsg,t-i}} = \frac{\left[\sum_{j \epsilon B_{dsgt}} \left(\frac{P_{jt}}{\varphi_{jt}} \right)^{1-\sigma_d^w} \right]^{\frac{1}{1-\sigma_d^w}}}{\left[\sum_{j \epsilon B_{dsg,t-i}} \left(\frac{P_{j,t-i}}{\varphi_{j,t-i}} \right)^{1-\sigma_d^w} \right]^{\frac{1}{1-\sigma_d^w}}}. \tag{5}$$

The price index for foreign products in store-category sg at time t is analogous.

Using the previous price index, the utility-maximizing quantity demanded of the domestic CES aggregate can be written as

$$Q_{dsgt} = \varphi_{dsgt} \left(\frac{P_{dsgt}}{P_{sgt}} \right)^{-\sigma_g^b} \left(\frac{Y_{sgt}}{P_{sgt}} \right)^{\epsilon_g^d} \tag{6}$$

where Y_{sgt} is expenditure in store-category sg at time t and P_{sgt} is an implicitly defined price index (Comin et al. (2015)). We have an analogous equation for the utility-maximizing quantity demanded of the foreign CES aggregate. In the next section we describe our empirical strategy for estimating our key structural parameters.

3 Empirical Strategy

In this section we discuss our empirical strategy for estimating the parameters in our model. We first structurally estimate the (micro) elasticity of substitution within domestic and foreign varieties in each category, using the identification via heteroskedasticity approach of Feenstra (1994). We then estimate the (macro) elasticity of substitution between domestic and foreign varieties in each category using store-category, share-weighted exchange rate index changes as an instrument for the change in the relative price index of foreign products. Therefore, we will estimate exchange rate pass-through before we estimate the macro elasticity.

3.1 Within elasticity

In applying Feenstra (1994) to our scanner data, we will make similar assumptions as those used in Broda and Weinstein (2010) and Hottman et al. (2016). First, we allow the marginal costs (C_{ut}) of each domestic variety $u \epsilon B_{dsgt}$ to be variable and increasing with output such that

$$C_{ut} = \delta_{ut}(1 + \omega_{dg}) Q_{ut}^{\omega_{dg}} \tag{7}$$

where $\omega_{dg} \geq 0$ determines the convexity of the cost function for domestic products in product category g and $\delta_{ut} > 0$ is a variety-level cost shifter[4]. Foreign varieties have the same cost structure, but their own convexity parameters ω_{fg} and variety-level cost shifters.

Barcode-level retail prices can be written as a markup μ_{gt} multiplied by the above marginal cost such that

$$P_{ut} = \mu_{gt} C_{ut} \tag{8}$$

We will be agnostic as to the source of the markup, as it could be set by retailers or manufacturers (or both, in the case of double-marginalization)[5]. Similarly, we also have a pricing equation for foreign products that follows the same structure.

[4]Equation 7 defines each supply curve in terms of the quantity supplied at the barcode-store level. For the majority of origin-product categories, our estimation results are little changed if we alternatively assume each supply curve is in terms of the quantity supplied of that barcode nationally.

[5]We will write the markup as μ_{gt} for simplicity, but we will implicitly allow the markup to vary by manufacturer-store-category-time when we implement estimation.

Next, we double difference the variety-level demand equation 3, where $\Delta^{k,t}$ is the time difference and the difference relative to the average variety k in the same store, product category, barcode manufacturer, and origin (domestic or foreign) as variety u:

$$\Delta^{k,t}\ln(S_{ut}) = (1-\sigma_d^w)\Delta^{k,t}\ln(P_{ut}) + e_{ut} \tag{9}$$

where $S_{ut} = (P_{ut}Q_{ut})/Y_{dsgt}$ is the share of domestic variety u among domestic products in store-group sg at time t (or analogously, the share among foreign varieties for a given foreign variety u). e_{ut} is a structural error term and can be expressed as

$$e_{ut} = (1-\sigma_d^w)\Delta^{k,t}\ln(\varphi_{ut}) \tag{10}$$

We also double-difference the pricing equation (in this case, for domestic goods) such that

$$\Delta^{k,t}\ln(P_{ut}) = \left(\frac{\omega_{dg}}{1+\omega_{dg}}\right)\Delta^{k,t}\ln(S_{ut}) + \kappa_{ut} \tag{11}$$

where κ_{ut} is also a structural error term that can be written as

$$\kappa_{ut} = \left(\frac{1}{1+\omega_{dg}}\right)\Delta^{k,t}\ln\delta_{ut} \tag{12}$$

Importantly, we assume that the demand and supply errors are orthogonal for each (in this case, domestic) variety u such that

$$G(\beta_{dg}) = \mathbb{E}_{\mathbb{T}}[x_{ut}(\beta_{dg})] = 0 \tag{13}$$

where $\beta_{dg} = \begin{bmatrix}\sigma_d^w \\ \omega_{dg}\end{bmatrix}$ and $x_{ut} = e_{ut}\kappa_{ut}$.

This orthogonality condition is plausible for a couple reasons. First, barcode characteristics are fixed over time. Second, variety and store-category-manufacturer-origin-time fixed effects have been differenced out, and thus the variation in prices due to markups has also been eliminated. Therefore, the supply shocks that remain are idiosyncratic shifts in the intercept of the supply curve. These shifts are likely not to be correlated with the remaining demand shocks that idosyncratically shift the intercept of the demand curve.

This orthogonality condition for each variety can be re-written as

$$\begin{aligned}\mathbb{E}_{\mathbb{T}}\left[\Delta\ln(P_{ut})^2\right] = &\left(\frac{1+\omega_{dg}(2-\sigma_d^w)}{(1-\sigma_d^w)(1+\omega_{dg})}\right)\mathbb{E}_{\mathbb{T}}\left[\Delta^{k,t}\ln(S_{ut})\Delta^{k,t}\ln(P_{ut})\right] \\ &-\left(\frac{\omega_{dg}}{(1+\omega_{dg})(1-\sigma_d^w)}\right)\mathbb{E}_{\mathbb{T}}\left[\Delta^{k,t}\ln(S_{ut})\Delta^{k,t}\ln(S_{ut})\right]\end{aligned} \tag{14}$$

Our orthogonality assumption is not itself sufficient for identification. However, if additionally $\mathbb{E}_{\mathbb{T}}[(\Delta^{k,t}\ln(P_{ut})^2]$, $\mathbb{E}_{\mathbb{T}}[\Delta^{k,t}\ln(S_{ut})\Delta^{k,t}\ln(P_{ut})]$, and $\mathbb{E}_{\mathbb{T}}[\Delta^{k,t}\ln(S_{ut})\Delta^{k,t}\ln(S_{ut})]$ are different across varieties, then we have N equations[6] and 2 unknown parameters (σ_d^w and ω_{dg}), which we can estimate using Generalized Method of Moments. We follow the same identification strategy for foreign varieties to obtain σ_f^w and ω_{fg}.

For each product category (and either foreign or domestic origin), stack the orthogonality conditions to form the GMM objective function

$$\hat{\beta}_{dg} = \arg\min_{\beta_{dg}} \left\{ G^*(\beta_{dg})' W G^*(\beta_{dg}) \right\}, \tag{15}$$

where $G^*(\beta_{dg})$ is the sample counterpart of $G(\beta_{dg})$ stacked over all barcodes in product category g for a given origin (in this case, domestic), and W is a positive definite weighting matrix. Following Broda and Weinstein (2010), we implement empirically a one-step GMM estimator that gives more weight to barcodes that are present in the data for longer time periods[7].

3.2 Between elasticity

After estimating our (micro) elasticities of substitution, σ_d^w and σ_f^w, we can construct the CES price indexes for foreign and domestic products in each store-category. We then double-difference our upper-tier demand equation 6 to get

$$\Delta^t \ln\left(\frac{Y_{fsgt}}{Y_{dsgt}}\right) = (1-\sigma_g^b)\Delta^t \ln\left(\frac{P_{fsgt}}{P_{dsgt}}\right) + (\epsilon_g^f - \epsilon_g^d)\Delta^t \ln\left(\frac{Y_{sgt}}{P_{sgt}}\right) + \varepsilon_{sgt} \tag{16}$$

where $\frac{Y_{fsgt}}{Y_{dsgt}}$ is spending on foreign products relative to domestic products in store-category sg at time t, $\frac{P_{fsgt}}{P_{dsgt}}$ is the price of foreign goods relative to domestic goods in store-category sg at time t, and $\frac{Y_{sgt}}{P_{sgt}}$ is real expenditure in store-category sg (nominal expenditure deflated by the implicit price index). We construct a store-category Törnqvist price index to use as a deflator in place of the unobserved, implicit price index.

Since the price of foreign goods relative to domestic goods is endogenous, we use as our main instrumental variable, the share-weighted exchange rate index for each store-category[8]. We construct this index using initial-period share weights of the bilateral exchange rates for each foreign country with products in the store-category. This is a store-level version of a trade-weighted exchange rate index. The four-quarter change in the share-weighted exchange rate index (X_{sgt}) in store-category sg at time t is constructed as

$$\Delta^t \ln X_{sgt} = \sum_{c \epsilon Csg} \left(\frac{Y_{csg,t=0}}{Y_{fsg,t=0}}\right) \Delta^t \ln E_{ct} \tag{17}$$

where $\frac{Y_{csg,t=0}}{Y_{fsg,t=0}}$ is the share of spending on products from foreign country c in the initial time

[6]Where N is the number of varieties in the origin-category.

[7]For GMM estimation, we also drop observations of barcodes that are not in the data for at least 20 quarters. We still use these barcodes to build subsequent price indexes and in all other results that follow.

[8]We also treat real category expenditure as endogenous, and use lagged nominal category expenditure as an additional instrument.

period for which category g exists in store s, and E_{ct} is country c's nominal bilateral exchange rate (foreign per dollar) at time t.

4 Data

Our retail scanner data comes from Information Resources Incorporated and contains barcode-level information on products in thousands of grocery and drug stores across the United States (Bronnenberg et al. (2008)). The data exists in weekly frequency from 2001 to 2012 and covers 30 different narrowly-defined product categories in 47 major markets. Examples of product categories are beer, cereal, shampoo, and toothpaste. For each barcode, we have information on product category, manufacturer, volume of package (e.g., oz of beer), and whether the product is a private label. For each store, we observe the weekly revenue and quantity sold by barcode. We define our barcode prices to be in terms of dollars per unit volume[9]. In their online appendix, Gagnon et al. (2017) show that category-level price indexes from this IRI data match the official BLS price indexes reasonably well.

Importantly, we have obtained novel barcode-level country of origin information from Label Insight, a vendor that digitizes product labels. Label Insight has information on over 80 percent of top-selling food, pet, and personal care items in the United States. This allows us to match our barcodes to a country of origin, which refers to the country in which the product was manufactured and packaged (including the United States). However, for our baseline analysis we also assume that products without origin information are U.S. goods. This is a reasonable assumption because imported final goods are required to have country of origin listed on the product label—for example, "Made in Japan" or "Product of Mexico"—but such requirements do not exist for products packaged and sold in the United States[10]. Therefore, U.S. products often do not specify a country of origin on the product label. On the other hand, the Federal Trade Commission regulates any "Made in the U.S.A" product claims, and requires for such a claim to be made that nearly all of the product's parts must originate in the United States[11]. In a robustness check in the Appendix A.1, we drop products of unknown origin from our analysis and show that our results are little changed.

We choose to focus our analysis on the ten narrow product categories in which the largest percentage of total sales are spent on foreign products (Table 1). These categories are beer, blades, cold cereal, deodorant, shampoo, soup, spaghetti sauce, sugar substitute, toothbrush, and toothpaste. We also drop drugstores and private-label barcodes from our data. Finally, we aggregate our variables to a quarterly frequency.

After limiting the analysis to ten categories, we have 41,709 unique barcodes across 2,503 stores. The products in these categories come from a total of 30 different countries, including the United States. Across all categories, 7.8 percent of total sales are spent on foreign origin

[9]That is, the price equals [revenue/(quantity*package volume)].

[10]Section 304, U.S. Tariff Act of 1930: "every article of foreign origin...imported into the United States shall be marked...to indicate to an ultimate purchaser in the United States the English name of the country of origin of the article".

[11]https://www.ftc.gov/tips-advice/business-center/guidance/complying-made-usa-standard

products and 92.2 percent are spent on U.S. products (including unknown origin products). Table 2 shows the top two foreign countries in terms of total sales on foreign origin products for each category[12]. As might be expected, most of the expenditures on foreign products are spent on goods originating in Canada and Mexico. However, expenditures on foreign blades, spaghetti sauce, and toothbrushes are concentrated on products from Brazil, the European Union, and China, respectively.

Table 1: Total share of spending on foreign barcodes

Product Category	2001Q1	2006Q1	2012Q1
beer	11.71%	14.69%	13.33%
blades	9.08%	7.51%	7.25%
cold cereal	2.35%	2.45%	6.18%
deodorant	0.70%	3.17%	4.85%
shampoo	0.46%	0.33%	1.00%
soup	0.49%	0.59%	5.06%
spaghetti sauce	0.98%	2.82%	4.94%
sugar substitute	0.0003%	0.008%	4.98%
toothbrush	1.08%	2.68%	8.39%
toothpaste	0.87%	0.84%	1.32%

Table 2: Top countries representing foreign sales

Product Category	% of Foreign Sales	
beer	Mexico (60%)	Netherlands (19%)
blades	Brazil (63%)	Germany (15%)
cold cereal	Mexico (56%)	Canada(43%)
deodorant	Canada (96%)	Thailand (2%)
shampoo	Canada (64%)	Italy (36%)
soup	Canada (79%)	Mexico (14%)
spaghetti sauce	Italy (99%)	Canada (.6%)
sugar substitute	Mexico (89%)	Italy (11%)
toothbrush	China (72%)	Germany (17%)
toothpaste	Mexico (88%)	Canada (12%)

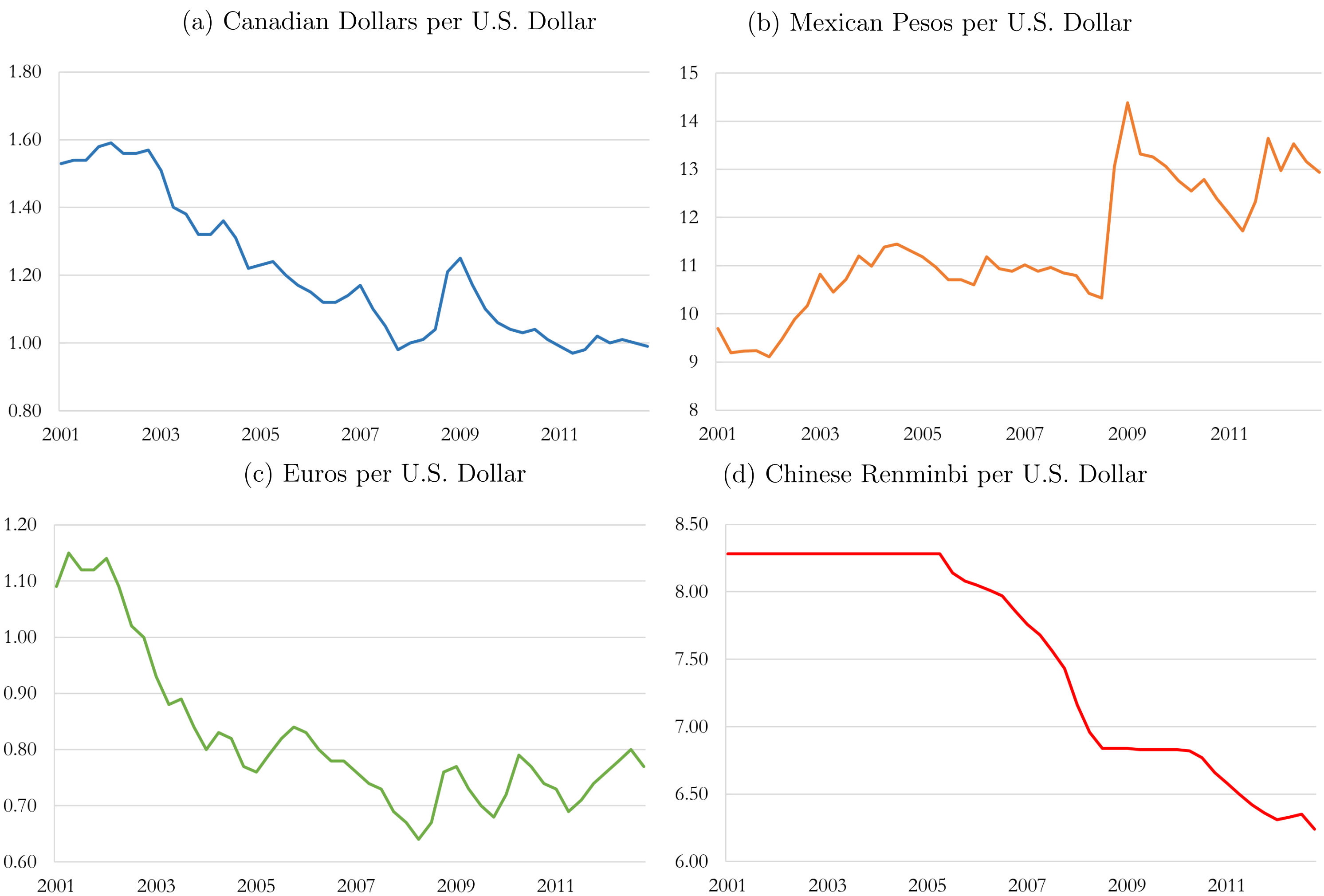
Percentages represent share of total foreign sales across all quarters.

We obtain bilateral exchange rates from Bloomberg at a quarterly frequency. Figure 1 shows the bilateral exchange rates between the U.S. dollar and the Canadian dollar, Mexican Peso, Euro, and Chinese Renminbi from 2001 to 2012. During this time period, the U.S. dollar is

[12]Among product categories for which we can easily find closely matching HS10 codes, we find a similar pattern of countries in average U.S. import data over 2001-2012. Blades is the main exception, in which the trade data show a larger role for China and a smaller role for Brazil.

largely depreciating against the Canadian dollar and the Euro, except for short periods of appreciation (in 2008 against the Canadian dollar and in 2008, 2009, and 2011 against the Euro). Prior to 2005, the Renminbi was fixed to the U.S. dollar. Since the exchange rate peg was lifted, the dollar has consistently depreciated against the Renminbi. By contrast, the dollar mostly appreciated against the Mexican Peso, particularly sharply in 2008 and in 2011.

Figure 1: Bilateral Exchange Rates



5 Estimation Results

5.1 Micro elasticities within foreign and domestic products

Table 3 shows our estimation results (with 95 percent confidence intervals) for the (micro) elasticity of substitution within domestic varieties, σ_d^w, and the elasticity of substitution within foreign varieties, σ_f^w, for the ten different product categories in our analysis.

Table 3: σ_d^w and σ_f^w estimates

Product Category	σ_d^w (95% C.I.)	N	R^2		σ_f^w (95% C.I.)	N	R^2
beer	6.50 (6.32, 6.67)	185,917	0.37		6.84 (6.61, 7.07)	24,822	0.44
blades	9.09 (8.65, 9.52)	33,722	0.22		5.68 (5.35, 6.01)	5,128	0.20
cold cereal	3.90 (3.88, 3.92)	197,447	0.54		4.19 (4.11, 4.27)	10,770	0.48
deodorant	5.61 (5.47, 5.75)	166,568	0.21		5.61 (5.23, 6.00)	6,808	0.33
shampoo*	5.44 (5.35, 5.52)	137,518	0.27		N/A	N/A	N/A
soup	4.07 (4.03, 4.10)	282,415	0.44		4.07 (3.96, 4.17)	13,738	0.46
spaghetti sauce	4.64 (4.52, 4.76)	62,493	0.43		3.11 (2.94, 3.27)	2,603	0.59
sugar substitute	5.00 (4.91, 5.09)	21,178	0.40		4.82 (4.32, 5.31)	250	0.64
toothbrush	5.75 (5.60, 5.89)	66,204	0.26		4.43 (4.00, 4.86)	2,798	0.38
toothpaste	4.96 (4.89, 5.03)	111,323	0.33		3.56 (3.27, 3.84)	387	0.69

*After dropping observations of barcodes that are not in the data for at least 20 quarters, the data does not contain enough observations for the estimation of σ_f^w for shampoo. However, we use σ_d^w as a substitute to construct the subsequent price index.

With these results, we can easily reject the hypothesis that σ_d^w and σ_f^w are statistically equal to 0 or 1, as our estimates and 95 percent confidence intervals are much larger. We therefore conclude that the demand for particular varieties within the set of domestic origin or foreign origin goods is elastic[13]. For example, our elasticity of substitution for domestic beer is 6.50. Therefore, if the price of one variety of domestic beer rises by 1 percent, we expect the quantity demanded for that variety to fall about 6.5 percent, relative to all other domestic beer varieties. Similarly, if the price of a foreign beer rises by 1 percent, the quantity demanded of that variety will fall about 6.84 percent, relative to other foreign beer varieties.

For the categories of beer, deodorant, soup, and sugar substitutes, σ_d^w and σ_f^w are very close to one another with overlapping confidence intervals, implying that foreign varieties are as substitutable with each other as domestic varieties are. By contrast, σ_d^w is significantly greater than σ_f^w for blades, spaghetti sauce, toothbrush, and toothpaste. Therefore, demand for domestic varieties in these categories is more responsive to price fluctuations than the demand for foreign varieties in the same categories. In other words, domestic varieties are more substitutable with one another.

5.2 Exchange rate pass-through

In this section, we report our estimates of exchange rate pass-through (which are also our first-stage estimation results for the macro elasticities of substitution between domestic and foreign varieties). This pass-through regression, in four-quarter changes, is given by

$$\Delta^t \ln\left(\frac{P_{fsgt}}{P_{dsgt}}\right) = \alpha_g + \beta_{gt} + \gamma_g \Delta^t \ln(X_{sgt}) + \delta_g \Delta^{t-1} \ln(Y_{sgt}) \tag{18}$$

[13] These estimated elasticities of substitution are comparable and broadly similar to those estimated on different scanner data in Hottman (2016), although that paper makes no distinction between domestic and foreign-origin products.

As mentioned in previous sections, $\frac{P_{fsgt}}{P_{dsgt}}$ is the price of foreign goods relative to domestic goods in store-category sg at time t, and X_{sgt} is a share-weighted bilateral exchange rate index based on the countries represented in store-category sg at time t. Additionally, $\Delta^{t-1}\ln(Y_{sgt})$ is the lagged change in total nominal expenditure in the store-category[14].

We are primarily interested in γ_g, the coefficient on exchange rates because this value can be interpreted as the exchange rate pass-through, or the effect of exchange rate changes on the prices of foreign goods relative to domestic goods. This coefficient, along with its 95 percent confidence interval, is reported in Table 4.

Table 4: First-stage results

Product Category	γ_g (95% C.I.)	N	F stat.
beer	-0.120 (-0.126, -0.114)	42,386	839.74
blades	-0.237 (-0.242, -0.232)	39,789	4241.40
cold cereal	-0.255 (-0.273, -0.237)	51,923	393.71
deodorant	-0.045 (-0.047, -0.042)	45,491	616.05
shampoo	-0.390 (-0.416, -0.363)	18,232	411.77
soup	-0.061 (-0.066, -0.056)	48,459	200.13
spaghetti sauce	-0.646 (-0.661, -0.631)	21,036	3584.85
sugar substitute	-0.346 (-0.417, -0.276)	1,209	46.74
toothbrush	-0.020 (-0.022, -0.019)	30,838	269.06
toothpaste	-0.374 (-0.384, -0.365)	36,287	3068.91

As expected, the exchange rate coefficient for each category is negative, and none of the confidence intervals contain 0. This means that as the dollar appreciates, the prices of foreign products will become cheaper relative to domestic products. Exchange rate pass-through varies significantly with product category, being the largest for spaghetti sauce (-0.65) and the smallest for toothbrushes (-0.02). Therefore, if the dollar appreciates 10 percent, the price of foreign spaghetti sauce will fall by 6.5 percent, relative to domestic spaghetti sauce. On the other hand, the relative price of foreign toothbrushes will only fall by 0.2 percent. For our median category, exchange rate pass-through is approximately -0.25 [15]. For comparison, we estimate exchange rate pass-through for beer to be about -0.12, which is similar to the Goldberg and Hellerstein (2012) estimate of -0.05 to -0.10.

5.3 Product Variety

The change in the relative CES price index for foreign goods can partly be driven by an expansion or contraction in product variety. For example, an appreciation of the dollar may increase foreign product variety by decreasing the prices of foreign products and encouraging greater imports.

[14] We include this additional instrument to deal with possible endogeneity in real category expenditure.

[15] Our pass-through (into retail prices) is a bit higher than estimates by Campa and Goldberg (2005), who find long-run pass-through of exchange rates into food import prices of about -0.20.

In this section, we assess the extent to which exchange rates affect foreign product variety, and the quantitative importance of this channel for our exchange rate pass-through estimates.

First, we regress the four-quarter percent change in the number of foreign barcodes against the four-quarter percent change in the share-weighted exchange rate index. This regression is given by

$$\Delta^t \ln(U^f_{sgt}) = \alpha_g + \beta_{gt} + \kappa_g \Delta^t \ln(X_{sgt}) \tag{19}$$

where U^f_{sgt} is the number of foreign barcodes in store-category sg at time t, and X_{sgt} is the share-weighted index of bilateral exchange rates in foreign currencies per dollar in store-category sg at time t.

The estimation results for κ_g and its 95 percent confidence intervals are reported in Table 5. The data indicate that there is a clear positive relationship between dollar appreciation and foreign product variety.

Table 5: Exchange Rate and Foreign Barcodes

Product Category	κ_g (95% C.I.)	N	R^2
beer	0.552 (0.548, 0.556)	42,386	0.66
blades	0.631 (0.627, 0.635)	39,789	0.78
cold cereal	0.877 (0.867, 0.887)	51,923	0.62
deodorant	0.224 (0.221, 0.227)	45,491	0.60
shampoo	0.753 (0.747, 0.758)	18,232	0.93
soup	0.317 (0.314, 0.321)	48,459	0.92
spaghetti sauce	0.953 (0.951, 0.955)	21,036	0.97
sugar substitute	0.175 (0.158, 0.192)	1,209	0.35
toothbrush	0.068 (0.066, 0.069)	30,838	0.40
toothpaste	0.614 (0.610, 0.618)	36,287	0.78

Since the CES price index reflects product entry and exit over time, our exchange rate pass-through results reflect this channel. To remove this entry and exit effect, we replace the CES index with an otherwise identical price index that eliminates the effects of product variety. Instead of Equation 5, we use

$$\frac{P_{dsgt}}{P_{dsg,t-4}} = \frac{\left[\sum_{j\epsilon(B_{dsgt}\bigcap B_{dsg,t-4})} \left(\frac{P_{jt}}{\varphi_{jt}}\right)^{1-\sigma^w_d}\right]^{\frac{1}{1-\sigma^w_d}}}{\left[\sum_{j\epsilon(B_{dsgt}\bigcap B_{dsg,t-4})} \left(\frac{P_{j,t-4}}{\varphi_{j,t-4}}\right)^{1-\sigma^w_d}\right]^{\frac{1}{1-\sigma^w_d}}} \tag{20}$$

There is an analogous equation for foreign products. This new price index change is constructed relative to four quarters prior and includes only products that exist in each store-category at time t and time $t-4$. Thus, we have eliminated product entry and exit.

Substituting this new price index for our original CES price index, we find that the exchange rate pass-through coefficient for the median category is approximately -0.11 (Table 6), about half of the original value of -0.25[16].

Table 6: First-stage results

Product Category	New γ_g (95% C.I.)	N	F stat.		Orig γ_g (95% C.I.)
beer	-0.081 (-0.090, -0.073)	42,377	196.98		-0.120 (-0.126, -0.114)
blades	-0.124 (-0.129, -0.118)	39,610	887.32		-0.237 (-0.242, -0.232)
cold cereal	-0.104 (-0.119, -0.089)	51,921	92.73		-0.255 (-0.273, -0.237)
deodorant	-0.013 (-0.016, -0.010)	45,265	54.09		-0.045 (-0.047, -0.042)
shampoo	-0.281 (-0.306, -0.255)	17,928	233.32		-0.390 (-0.416, -0.363)
soup	-0.024 (-0.029, -0.019)	48,421	49.92		-0.061 (-0.066, -0.056)
spaghetti sauce	-0.278 (-0.294, -0.261)	20,654	541.33		-0.646 (-0.661, -0.631)
sugar substitute	-0.166 (-0.213, -0.118)	1,164	25.74		-0.346 (-0.417, -0.276)
toothbrush	-0.003 (-0.005, -0.001)	30,679	17.12		-0.020 (-0.022, -0.019)
toothpaste	-0.117 (-0.123, -0.110)	35,452	634.76		-0.374 (-0.384, -0.365)

5.4 Macro elasticities between domestic and foreign products

Following our first-state estimation results from Equation 18, we can then run the second stage regression. This empirical version of Equation 16 is given by

$$\Delta^t \ln\left(\frac{Y_{fsgt}}{Y_{dsgt}}\right) = \alpha_g + \beta_{gt} + (1-\sigma_g^b)\Delta^t \widehat{\ln\left(\frac{P_{fsgt}}{P_{dsgt}}\right)} + (\epsilon_g^f - \epsilon_g^d)\Delta^t \widehat{\ln\left(\frac{Y_{sgt}}{P_{sgt}}\right)} + \varepsilon_{sgt} \tag{21}$$

We estimate the (macro) elasticities of substitution between domestic and foreign varieties, σ_g^b, along with the coefficient on income, $(\epsilon_g^f - \epsilon_g^d)$ for each product category. These values and their 95 percent confidence intervals are reported in Table 7.

[16]For comparison, Nakamura and Steinsson (2012) estimate that not accounting for product replacement bias results in an exchange rate pass-through coefficient for U.S. import prices that is about one-third to one-half the size of the coefficient when accounting for the bias.

Table 7: Second-stage results

Product Category	σ_g^b (95% C.I.)	$(\epsilon_g^f - \epsilon_g^d)$ (95% C.I.)	N	F stat.
beer	2.84 (2.57, 3.11)	-0.004 (-0.006, -0.0008)	42,386	90.14
blades	2.97 (2.89, 3.06)	-0.005 (-0.010, 0.001)	39,789	1014.89
cold cereal	2.66 (2.44, 2.88)	-0.003 (-0.005, -0.0001)	51,923	108.35
deodorant	3.93 (3.55, 4.31)	-0.013 (-0.028, 0.001)	45,491	153.13
shampoo	2.52 (2.33, 2.71)	-0.007 (-0.025, 0.010)	18,232	273.29
soup	2.80 (2.64, 2.97)	-0.001 (-0.006, 0.005)	48,459	229.35
spaghetti sauce	2.18 (2.07, 2.29)	-0.025 (-0.088, 0.038)	21,036	625.79
sugar substitute	1.40 (1.18, 1.61)	0.004 (-0.052, 0.060)	1,209	34.09
toothbrush	3.81 (3.42, 4.19)	0.005 (-0.001, 0.011)	30,838	110.09
toothpaste	2.55 (2.33, 2.78)	0.032 (-0.112, 0.175)	36,287	452.58

As with the estimates for (micro) elasticities of substitution within foreign and domestic varieties, σ_d^w and σ_f^w, we can also reject the hypothesis that our macro elasticities, σ_g^b, are 0 or 1. Estimates and confidence intervals for σ_g^b are greater than one, suggesting that the demands for foreign and domestic varieties as a whole are elastic[17]. For example, if the price of domestic beer rises by 1 percent, then the real consumption of foreign beer relative to domestic beer will rise by 2.84 percent. When comparing σ_g^b to σ_d^w and σ_f^w, we find that σ_g^b is smaller for every category-origin, with the only exception being toothbrushes where the confidence interval for σ_g^b overlaps with the confidence interval for σ_f^w. Therefore, we find clear evidence that domestic varieties are more substitutable with one another than they are with imported products. Analogously, foreign varieties are also more substitutable with one another than they are with domestic products.

We find minimal effects of non-homothetic expenditure switching between foreign and domestic products. Most of our real expenditure coefficients are both economically and statistically insignificant. However, we find statistically significant results of expenditure switching for beer and cold cereal. The slightly negative expenditure coefficients for these two categories indicate that as real expenditure in the store increases, consumption switches away from foreign products and towards domestic products. This suggests that foreign beer and cereal could be of a slightly lower quality.

5.5 Model fit

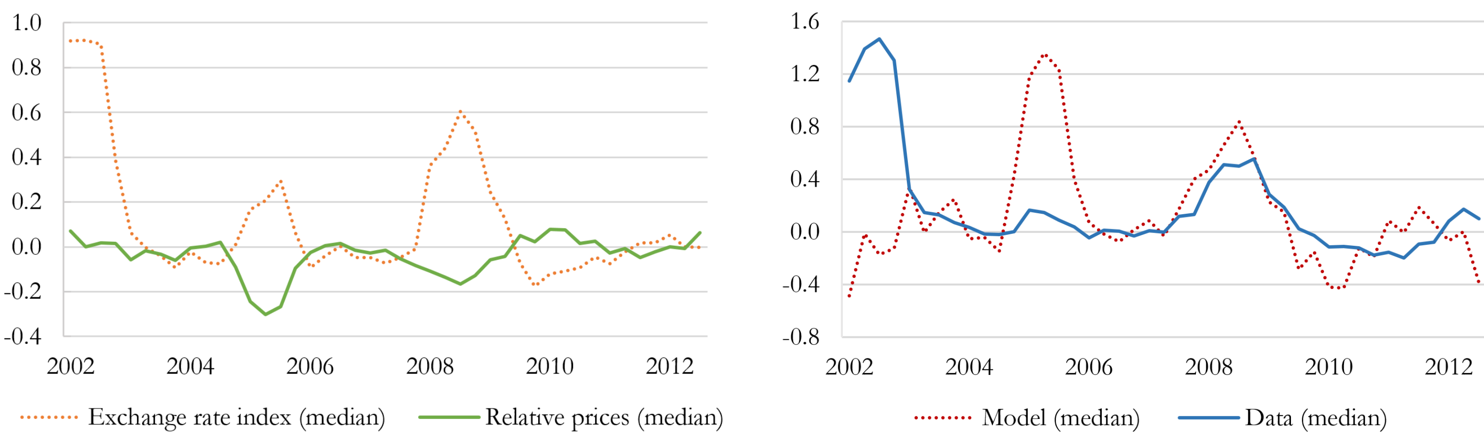
This section provides examples of the relationship between relative prices and exchange rate indexes in our data, as well as a graphical assessment of the fit of our model of expenditure switching (the macro elasticity). First, the left hand side of Figure 2 shows four-quarter log-changes in the exchange rate index and the relative price of foreign to domestic products for the median store in the deodorant product category. This figure shows that (with the exception of the 2002 period) when the exchange rate index rises as the dollar appreciates, prices of foreign

[17]For rough comparison, Feenstra et al. (2018) estimate a range of macro elasticities between 2.27 and 4.08 for their food products category.

deodorants tend to fall relative to domestic ones. The right hand side of Figure 2 plots the model predicted four-quarter log-changes in the relative expenditure on foreign products, based on only the realized path of relative prices and real store-category expenditure[18], for the median store in the deodorant category. The model misses the relative growth in spending on foreign products in the 2002 period (as relative prices were little changed in this period), and overpredicts the growth in spending on foreign products in the 2005 period. However, the model matches the data very well in the second half of the sample period.

Figure 2: Deodorant

(a) Prices vs. Exchange Rates (4-quarter Δ) (b) Foreign vs. U.S. Goods Sales (4-quarter Δ)

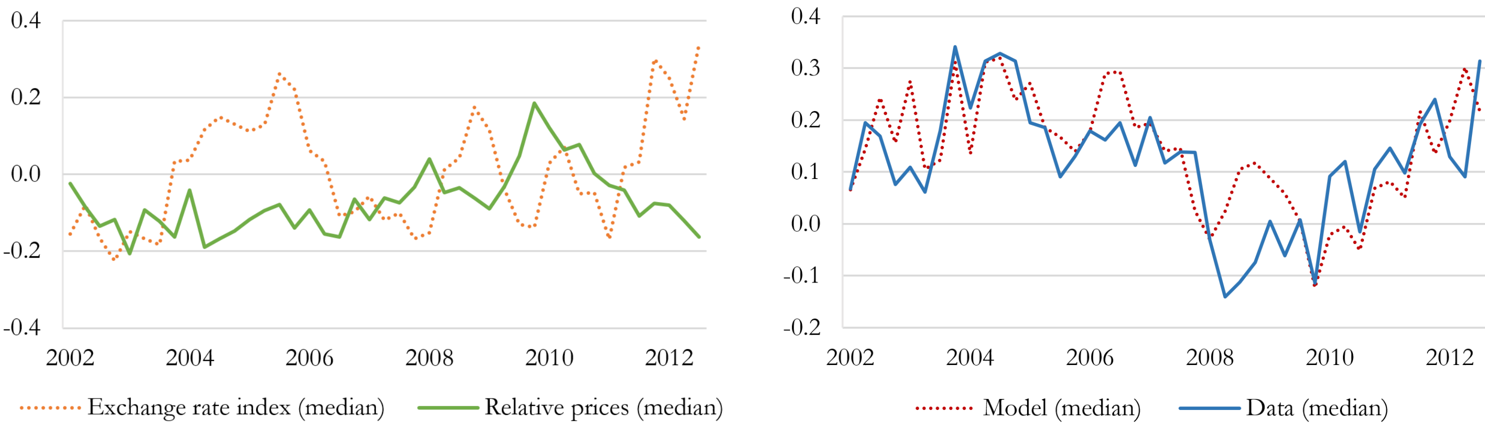


As another example, the left hand side of Figure 3 shows that for the median store in the spaghetti sauce category, the relative prices of foreign products tends to move inversely with the exchange rate index. The right hand side of Figure 3 shows that for the median store, the model prediction fits the data well over the full sample period.

[18] We are not using the model's taste shifters, which rationalize residual expenditure changes.

Figure 3: Spaghetti Sauce

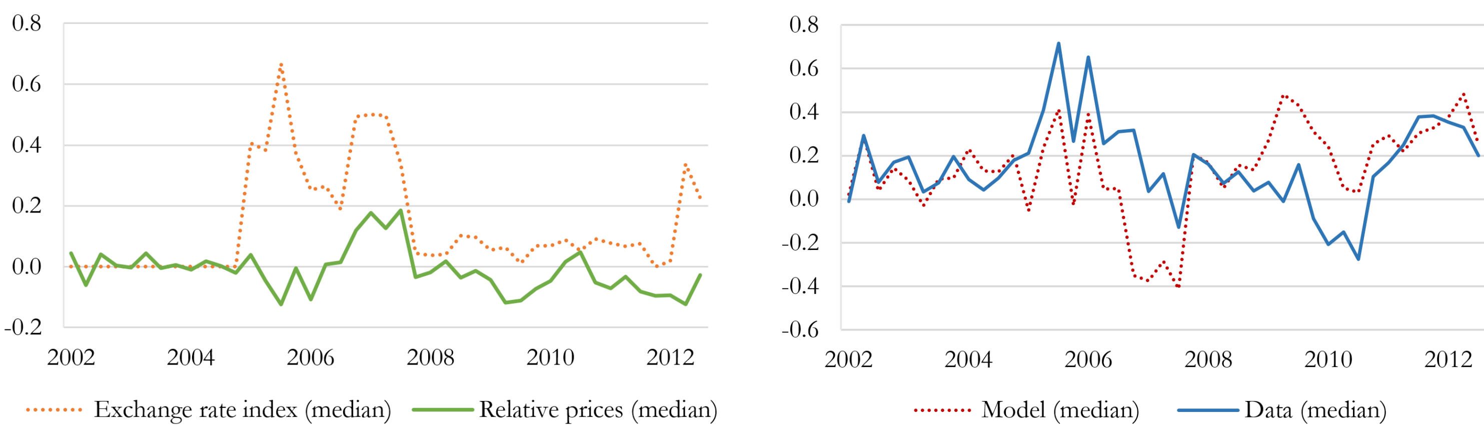
(a) Prices vs. Exchange Rates (4-quarter Δ) (b) Foreign vs. U.S. Goods Sales (4-quarter Δ)



The left hand side of Figure 4 shows the relative prices of foreign products and the exchange rate index for the median store in the toothbrush category. Relative foreign prices appear to decrease when the dollar appreciates, but the response is not particularly strong in the 2006 period. The right hand side of Figure 4 shows that for the median store, the model prediction fits the data reasonably well over the sample period. The main model misses are an under-prediction of foreign spending growth in 2007 and and over-prediction of foreign spending growth in 2009.

Figure 4: Toothbrush

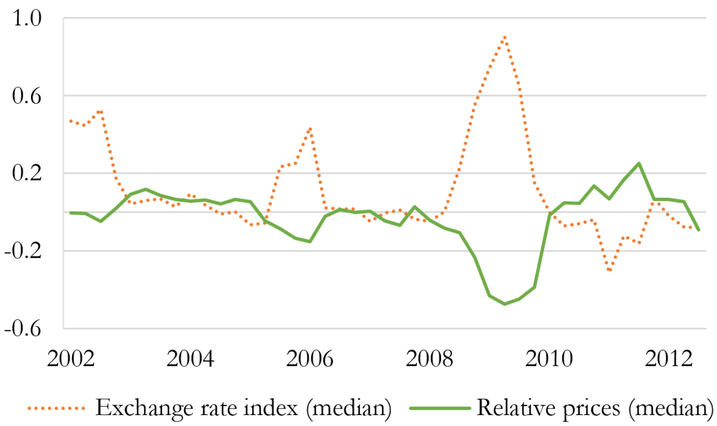
(a) Prices vs. Exchange Rates (4-quarter Δ) (b) Foreign vs. U.S. Goods Sales (4-quarter Δ)



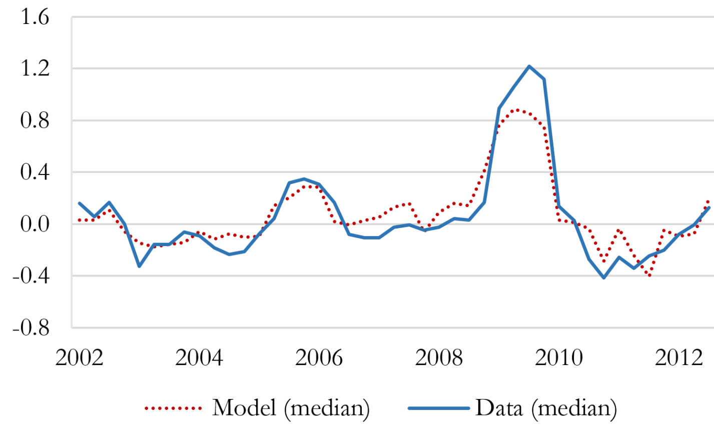
Finally, the left hand side of Figure 5 shows that for the median store in the toothpaste category, the relative price of foreign products tends to clearly move inversely with the exchange rate index. The right hand side of Figure 5 shows that the model prediction fits the data, for the median store in the toothpaste category, extremely well over the full sample period.

Figure 5: Toothpaste

(a) Prices vs. Exchange Rates (4-quarter Δ)



(b) Foreign vs. U.S. Goods Sales (4-quarter Δ)



6 Conclusion

This paper is the first to combine U.S. retail scanner data with barcode-level country of origin data. We use this novel dataset for 10 narrow product categories from 2001 to 2012 to estimate a nested, non-homothetic demand system. Our results shed light on several related questions.

We estimate parameters for the (micro) elasticities of substitution within domestic and foreign products separately in each product category. For about half of our product categories, the within elasticity for foreign products is smaller than the elasticity for domestic products. For the other half of our categories, the within elasticity parameters are about the same for foreign and domestic products.

Using CES price indexes, we estimate exchange rate pass-through of about -0.25 for the median product category. We also document a strong correlation between the number of foreign products on store shelves and the exchange rate. When we re-estimate exchange rate pass-through after excluding the effect of product entry and exit from the CES price index, we find that exchange rate pass-through is about half its previous magnitude.

Finally, we use the changes in the relative price of foreign to domestic products that result from exchange rate changes to estimate the (macro) elasticity of substitution between foreign and domestic products in each product category. We find that the between elasticity is significantly smaller than the within elasticity for essentially all product categories. However, we find little evidence of income-induced expenditure switching.

Therefore, our results overall provide uniquely clean evidence that support the views of Nakamura and Steinsson (2012) and Feenstra et al. (2018), while challenging the applicability of the results of Bems and Di Giovanni (2016) to a country like the United States.

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Appendix

A.1 Eliminating unknown-origin products

In our baseline analysis, we make the assumption that products for which we do not have country of origin information are domestic products. This is because items made and packaged in the United States do not need to have "United States" or "USA" mentioned on the packaging. Although we think this assumption is reasonable, we conduct a robustness check on our results by eliminating all products of unknown origin. Once dropping these 38,268 unique barcodes of unknown origin (making up 94.1 percent of total barcodes and 77.6 percent of total sales in these ten categories), we end with 2,441 unique barcodes for which we definitely know country of origin (including "Made in the U.S.A."). About two-thirds of total sales on remaining barcodes are spent on domestic goods, and about one-third are spent on foreign goods.

We report below the results from this robustness check. Table 8 contains estimates for the new σ_d^w, along with the original σ_d^w for comparison. Table 9 contains the new and original estimates for exchange rate pass-through. Finally, Table 10 contains the new and original estimates of σ_g^b.

Table 8: σ_d^w estimates

Product Category	New σ_d^w (95% C.I.)	N	R^2		Orig σ_d^w (95% C.I.)
beer	8.33 (7.94, 8.73)	12,095	0.22		6.50 (6.32, 6.67)
blades*	N/A	N/A	N/A		9.09 (8.65, 9.52)
cold cereal	4.56 (4.29, 4.83)	9,242	0.54		3.90 (3.88, 3.92)
deodorant	5.17 (4.95, 5.38)	20,744	0.37		5.61 (5.47, 5.75)
shampoo	4.03 (3.94, 4.13)	34,046	0.37		5.44 (5.35, 5.52)
soup	3.89 (3.86, 3.93)	175,992	0.40		4.07 (4.03, 4.10)
spaghetti sauce	3.34 (3.13, 3.56)	8,632	0.37		4.64 (4.52, 4.76)
sugar substitute*	N/A	N/A	N/A		5.00 (4.91, 5.09)
toothbrush*	N/A	N/A	N/A		5.75 (5.60, 5.89)
toothpaste	4.63 (4.31, 4.94)	3,845	0.52		4.96 (4.89, 5.03)

*After dropping products of unknown origin, the data does not contain enough observations for the estimation of σ_d^w for blades, sugar substitutes, and toothbrushes. However, we use σ_f^w as a substitute to construct the subsequent price indexes.

Table 9: First-stage results

Product Category	New γ_g (95% C.I.)	N	F stat.		Orig γ_g (95% C.I.)
beer	-0.123 (-0.130, -0.116)	42,332	574.83		-0.120 (-0.126, -0.114)
cold cereal	-0.281 (-0.313, -0.249)	48,646	148.84		-0.255 (-0.273, -0.237)
deodorant	-0.043 (-0.046, -0.041)	45,490	503.82		-0.045 (-0.047, -0.042)
shampoo	-0.479 (-0.507, -0.451)	18,232	559.05		-0.390 (-0.416, -0.363)
soup	-0.057 (-0.062, -0.052)	48,459	238.99		-0.061 (-0.066, -0.056)
spaghetti sauce	-0.619 (-0.635, -0.603)	20,965	2767.15		-0.646 (-0.661, -0.631)
sugar substitute	-0.357 (-0.430, -0.284)	1,162	46.42		-0.346 (-0.417, -0.276)
toothbrush	-0.024 (-0.027, -0.020)	28,616	112.90		-0.020 (-0.022, -0.019)
toothpaste	-0.380 (-0.390, -0.370)	36,251	2992.17		-0.374 (-0.384, -0.365)

*After dropping products of unknown origin, the data does not contain enough observations to evaluate the exchange rate pass-through onto blades' prices.

Table 10: Second-stage results

Product Category	New σ_g^b (95% C.I.)	N	F stat.		Orig σ_g^b (95% C.I.)
beer	2.64 (2.37, 2.91)	42,332	70.07		2.84 (2.57, 3.11)
cold cereal	2.33 (2.04, 2.62)	48,646	40.50		2.66 (2.44, 2.88)
deodorant	3.95 (3.59, 4.31)	45,490	167.64		3.93 (3.55, 4.31)
shampoo	2.21 (2.10, 2.33)	18,232	405.96		2.52 (2.33, 2.71)
soup	2.90 (2.71, 3.08)	48,459	203.23		2.80 (2.64, 2.97)
spaghetti sauce	2.23 (2.05, 2.41)	20,965	343.78		2.18 (2.07, 2.29)
sugar substitute	1.44 (1.26, 1.63)	1,162	31.16		1.40 (1.18, 1.61)
toothbrush	4.30 (3.77, 4.84)	28,616	77.74		3.81 (3.42, 4.19)
toothpaste	2.53 (2.32, 2.74)	36,251	497.75		2.55 (2.33, 2.78)

After dropping products of unknown origin, the data does not contain enough observations to evaluate σ_g^b for blades.

Upon dropping all products of unknown origin, our results remain very similar to those in Section 6. Our exchange-rate pass through coefficients for each category are nearly identical to the original coefficients. We estimate exchange rate pass-through to be -0.28 for the median category, similar to our original estimate of -0.25. Again, in all of the categories for which we obtain results, our estimates for σ_g^b are nearly identical to the original coefficients. Finally, for all categories, estimates for σ_g^b are still less than our estimates for σ_d^w and σ_f^w, except for in the toothbrush category in which σ_g^b still overlaps with σ_f^w.