Web Appendix: Income-Induced Expenditure Switching*

Published in the *American Economic Review*, December 2016, 106(12): 3989–3931

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Abstract

This paper shows that an income effect can drive expenditure switching between domestic and imported goods. We use a unique Latvian scanner-level dataset, covering the 2008–09 crisis, to document several empirical findings. First, expenditure switching accounted for one-third of the fall in imports, and took place *within* narrowly-defined product groups. Second, there was no corresponding within-group change in relative prices. Third, consumers substituted from expensive imports to cheaper domestic alternatives. These findings motivate us to estimate a model of non-homothetic consumer demand, which explains two-thirds of the observed expenditure switching. Estimated switching is driven by income, not changes in relative prices.

JEL Classifications: F1; F3; F4

Keywords: Expenditure switching; relative price adjustment; crisis; income effect

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Appendix A Role of Intensive and Extensive Margins in Expenditure Switching

Given that we are using detailed item level data, we wish to investigate the potential impact of entry and exit of items on the dynamics of expenditures, both for domestic and foreign goods. There are two important reasons to do so. First, as recently shown by Corsetti, Martin and Pesenti (2013), it is theoretically possible to have expenditure switching without a corresponding relative price change if there is substantial entry and exit of goods. Second, our modeling and estimation strategies in Section 4 rely on continuing items as the source of identification.

In order to examine the importance of entry and exit in our data, we follow two different strategies. First, we consider a gross concept, and look at the time series of items, aggregated by their domestic/foreign origin, for continuing, entering and exiting items. Figure A1 plots these time series for q-o-q data. The top panel graphs the count of UPC items, while the bottom panel plots the time series based on total expenditures. Regardless of the measure, continuing items make up the largest component of total of goods, both for domestic and foreign items, over time. Moreover, in terms of expenditures, continuing items capture the boom-bust cycle as well as the expenditure switching from imported to domestic items during the crisis.

Second, to more directly examine the role of entering/exiting versus continuing items in expenditure switching, we decompose the *growth rate* of expenditure switching into contributions from intensive and extensive margins. Borrowing from the methodology di Giovanni, Levchenko and Méjean (2014), we decompose a growth rate of a given variable, which is constructed using item (i) and product group (g) data. In particular, for simplicity we will consider the growth rate of total sales, X_t, which are the sum of individual item sales, x_{igt}, where an item i falls into a group g. We will consider the growth rate between $t-1$ and t.

The log-difference growth rate of total sales can be manipulated to obtain an (exact) decomposition into intensive and extensive components:

$$\begin{aligned}\tilde{\gamma}_t &\equiv \ln \sum_{i\in I_t} x_{igt} - \ln \sum_{i\in I_{t-1}} x_{igt-1} \\ &= \ln \frac{\sum_{i\in I_{t/t-1}} x_{igt}}{\sum_{i\in I_{t/t-1}} x_{igt-1}} - \left(\ln \frac{\sum_{i\in I_{t/t-1}} x_{igt}}{\sum_{i\in I_t} x_{igt}} - \ln \frac{\sum_{i\in I_{t/t-1}} x_{igt-1}}{\sum_{i\in I_{t-1}} x_{igt-1}} \right) \\ &= \underbrace{\gamma_t}_{\text{Intensive margin}} - \underbrace{\ln \frac{\pi_{t,t}}{\pi_{t,t-1}}}_{\text{Extensive margin}} ,\end{aligned} \tag{A.1}$$

where $I_{t/t-1}$ is the set of items sold in both t and $t-1$ (the intensive sub-sample of items in year t) and $\pi_{t,t}$ ($\pi_{t,t-1}$) is the share of items sold in this intensive sub-sample of goods in period t ($t-1$). Entrants have a positive impact on growth while exiters push the growth rate down, and the net impact is proportional to the share of entrants'/exiters' sales in aggregate sales.[1] Meanwhile, an observation only belongs to the intensive margin if an individual firm serves an individual destination in both periods.

The growth rate decomposition of total sales, (A.1), can be arbitrarily applied to total sales, total import sales, or total domestic sales in Latvia. This is the crucial point to consider when calculating the decomposition for the growth rate of expenditure switching. Let us define the share of imported items to total items for the overall economy at t, s_t^F as:

$$s_t^F = \frac{X_t^F}{X_t}, \tag{A.2}$$

where X_t^F are total imports at t. Then the (log) growth rate of s_t^F – i.e., the growth rate of *expenditure switching* – can be defined as a function of the growth rate of imports and total sales:

$$\begin{aligned} \ln s_t^F &= \ln X_t^F - \ln X_t \\ &= \ln \sum_{i \in I_t} x_{igt}^F - \ln \sum_{i \in I_t} x_{igt}. \end{aligned}$$

Therefore, the growth rate of s_t^F between $t-1$ and t is:

$$\begin{aligned} \ln s_t^F - \ln s_{t-1}^F &= \left(\ln \sum_{i \in I_t} x_{igt}^F - \ln \sum_{i \in I_t} x_{igt} \right) - \left(\ln \sum_{i \in I_{t-1}} x_{igt-1}^F - \ln \sum_{i \in I_{t-1}} x_{igt-1} \right) \\ &= \left(\ln \sum_{i \in I_t} x_{igt}^F - \ln \sum_{i \in I_{t-1}} x_{igt-1}^F \right) - \left(\ln \sum_{i \in I_t} x_{igt} - \ln \sum_{i \in I_{t-1}} x_{igt-1} \right) \\ &= \tilde{\gamma}_t^F - \tilde{\gamma}_t. \end{aligned} \tag{A.3}$$

We can therefore apply the decomposition (A.1) to the total growth rate of imports ($\tilde{\gamma}_{At}^F$) and total sales ($\tilde{\gamma}_{At}$), and take their difference to obtain an exact decomposition of the intensive and extensive components of the growth rate of expenditure switching over time.

We calculate the overall, intensive and extensive growth rates from q-o-q growth in expenditure switching and then sum the growth rates over a four-quarter overlapping rolling window, in order to avoid seasonality. Figure A2 plots the results. First, the import

[1] This decomposition follows the same logic as the decomposition of price indices proposed by Feenstra (1994).

expenditure share fell by around 10% during the crisis. Since imports account for slight more than 1/3 of expenditures, this fall in imports is consistent with a 3.8% of expenditures allocated towards domestic items, Second, the intensive component tracks very closely the growth rate of the aggregate expenditure switching during the crisis. Third, the growth rate of the extensive component during the crisis is relatively flat and positive, indicating a small but persistent switching of expenditures towards imported rather than domestic items. All in all, this decomposition assuages our concern that ignoring the extensive margin in our analysis will lead to any misleading conclusions.

Appendix B Decomposition of the Within Expenditure Switching: Within/Across Store Components

This appendix further decomposes expenditure switching within product groups into switching within/across store types. Expenditure switching within a product group g can be expressed as

$$\varphi^F_{gt} - \varphi^F_{gk} = \textstyle\sum_v m_{gvt}\varphi^F_{gvt} - \sum_v m_{gvk}\varphi^F_{gvk},$$

where $v = \{H, S, D\}$ indexes the three store types in our dataset, m_{gvt} is the share of store v in total expenditures on group g in period t and φ^F_{gvk} is the share of imports in product group g and store v at time t.

We can then decompose expenditure switching within a product group g as

$$\varphi^F_{gt} - \varphi^F_{gk} = \underbrace{\textstyle\sum_v m_{gvk}\left(\varphi^F_{gvt} - \varphi^F_{gvk}\right)}_{\text{Within stores}} + \underbrace{\textstyle\sum_v \varphi^F_{gvk}\left(m_{gvt} - m_{gvk}\right)}_{\text{Across stores}} + \textstyle\sum_v \Delta\varphi^F_{gvt}\Delta m_{gvt}.$$

Aggregate expenditure switching within product groups, as defined by the first term on the right hand side of equation (1), can then be decomposed into within and across store types as

$$\underbrace{\textstyle\sum_g s_{gk}\left(\varphi^F_{gt} - \varphi^F_{gk}\right)}_{\text{Within groups}} = \underbrace{\textstyle\sum_g s_{gk}\sum_v m_{gvk}\left(\varphi^F_{gvt} - \varphi^F_{gvk}\right)}_{\text{Within stores}} + \underbrace{\textstyle\sum_g s_{gk}\sum_v \varphi^F_{gvk}\left(m_{gvt} - m_{gvk}\right)}_{\text{Across stores}} + \underbrace{\textstyle\sum_g s_{gk}\sum_v \Delta\varphi^F_{gvt}\Delta m_{gvt}}_{\approx 0}.$$

The above equation decomposes the overall within contribution to expenditure switching into two subcomponents: (i) within groups and within a store type, and (ii) within groups, but across store types. Similar to the decomposition of aggregate expenditure switching in equation (1), the within/within margin contributes directly to within expenditure switching,

as consumers substitute between domestic and imported goods within a product group and a particular store type. Within/across expenditure switching can contribute indirectly if consumers reallocate expenditures across stores and groups' import shares across stores differ.

The decomposition results are reported in Figure A3 and show that the within expenditure switching took place almost entirely within store types. We further find that switching within each of the three store types contributed similarly to the overall expenditure switching. The across store component contributed less than 10% to the overall switching within 4-digit product groups.

This finding is not surprising when interpreted in terms of savings that consumers could make by switching across items within a product group in a given store as opposed to switching across stores. Specifically, by comparing prices across store types, we find that in Supermarkets and Hypermarkets 70% of monthly prices of overlapping UPCs items are identical, and in 97% of cases the deviation in prices is less than 5%. The mean item price in Supermarkets is only 0.07% below the corresponding price in the Hypermarket. Prices in Discounter stores are on average 12.7% lower than in Supermarkets and Hypermarkets, while the median UPC item is 10.6% cheaper. These price differentials imply a small margin for savings when compared to the within group/within store item price dispersion, which we discuss in Section 3.3.

Looking further into the contributing factors to the limited overall switching across store types, we find some systematic differences in import shares across stores. Aggregate imports shares for Discounter stores, Supermarkets and Hypermarkets are 0.30, 0.41 and 0.50, respectively. However, shares of the three store types in total expenditures by product groups did not vary systematically during the crisis.

Appendix C Demand Model and Estimation Derivation

C.1 Setup

Define the expenditure allocation problem over F&B for a representative consumer as

$$\max_{\{c_{igt}\}} U_t = \left(\sum_g \omega_g^{\frac{1}{\rho}} c_{gt}^{\frac{\rho-1}{\rho}} \right)^{\frac{\rho}{\rho-1}}$$

$$c_{gt} = \left(\left(\frac{1}{N_{gt}} \right)^{\frac{1}{\sigma_g}} \sum_{i \in I_{gt}} \widehat{c}_{igt}^{\frac{\sigma_g - 1}{\sigma_g}} \right)^{\frac{\sigma_g}{\sigma_g - 1}}, \text{ where } \widehat{c}_{igt} = \theta_{ig}^{\lambda_g(C_t)} c_{igt}$$

s.t.

$$\sum_g \sum_i p_{igt} c_{igt} = C_t.$$

Utility is defined over G product groups with the familiar CES aggregator. Within each product group g a consumer chooses between a group-specific set of items (there are N_{gt} items), each denoted $\widehat{c}_{igt}$, measured in 'utils,' and constructed as $\widehat{c}_{igt} = \theta_{ig}^{\lambda_g(C_t)} c_{igt}$, where c_{igt} is measured in common physical units (e.g., KG or L) and θ_{ig} is a factor that converts physical units into 'utils.' In Hallak (2006), θ_{ig} is as a proxy for quality differences and is measured using export unit values. We follow the same strategy using the UPC-level unit values, though as discussed above, there might be other factors driving the difference in unit values than just quality. Furthermore, as in Hallak (2006), we allow θ_{ig} to vary with income level (measured as total expenditures C_t), so that the degree to which "quality differences" within a product group matter is an increasing function of income. Specifically, $\lambda_g(C_t)$ captures the consumer's intensity for demand of an item's "quality" in a given group g, and varies with income C_t such that $\partial \lambda_g(C_t)/\partial C_t > 0$. It is worth stressing again that the specified model does not differentiate between domestic and foreign goods within a product group. We also allow for the elasticity of substitution between items within a group, σ_g, and the number of items within a group N_g, to vary by product group.

C.2 Characterization of the Model Solution

Given prices, p_{igt}, total expenditure, C_t, qualities, θ_{ig}, and parameter values, the consumer optimally allocates food expenditures in each period. Because modifications to the standard CES utility function rely entirely on exogenous parameters, the familiar first-order conditions hold both at the top and bottom levels of the utility. Specifically, at the top level we have

$$c_{gt} = \omega_g P_{gt}^{-\rho} C_t,$$

and consistent with the expenditure share notation in the previous section, group g's expenditure share can be written as

$$s_{gt} \equiv \frac{P_{gt} c_{gt}}{C_t} = \omega_g P_{gt}^{1-\rho}. \tag{C.1}$$

The utility-based aggregate price index, which we use as a numéraire, is

$$P_t = \left(\sum_g \omega_g P_{gt}^{1-\rho} \right)^{\frac{1}{1-\rho}}.$$

At the bottom level of the utility, i.e., within product groups, the demand equation is

$$c_{igt} = \frac{1}{N_{gt}\theta_{ig}^{\lambda_g(C_t)}} \left(\frac{\frac{p_{igt}}{\theta_{ig}^{\lambda_g(C_t)}}}{P_{gt}} \right)^{-\sigma_g} c_{gt},$$

so that an item's within-group expenditure share is

$$\varphi_{igt} \equiv \frac{p_{igt}c_{igt}}{P_{gt}c_{gt}} = \frac{1}{N_{gt}} \left(\frac{\frac{p_{igt}}{\theta_{ig}^{\lambda_g(C_t)}}}{P_{gt}} \right)^{1-\sigma_g}, \tag{C.2}$$

and the item's expenditure share in total F&B expenditures is

$$s_{igt} \equiv \varphi_{igt}s_{gt} = \frac{1}{N_{gt}} \left(\frac{\frac{p_{igt}}{\theta_{ig}^{\lambda_g(C_t)}}}{P_{gt}} \right)^{1-\sigma_g} \omega_g P_{gt}^{1-\rho}. \tag{C.3}$$

Finally, the utility-based price index for a product group is

$$P_{gt} = \left(\frac{1}{N_{gt}} \sum_i \left(\frac{p_{igt}}{\theta_{ig}^{\lambda_g(C_t)}} \right)^{1-\sigma_g} \right)^{\frac{1}{1-\sigma_g}}. \tag{C.4}$$

It is instructive to note that if the income level and quality considerations are switched off, i.e., $\lambda_g(C_t) = 0$, then the equation for s_{igt} collapses to

$$s_{igt} = \frac{1}{N_{gt}} \left(\frac{p_{igt}}{P_{gt}} \right)^{1-\sigma_g} \omega_g P_{gt}^{1-\rho},$$

which is the standard CES expression for the item's expenditure share in total expenditures. However, more generally income affects the expenditure share, so that the demand system is non-homothetic.

Equilibrium: *Given prices, p_{igt}, total expenditure, C_t, qualities, θ_{ig}, and parameter values, a consumer optimally allocates food expenditures in each period. The solution of the demand system can be characterized by a system of expenditure share equations s_{igt}, combined with group and aggregate price indexes and the budget constraint. One can solve the system to obtain the optimal consumption quantities for each item, c_{igt}.*

C.3 Estimation Equation

The key equation that characterizes the solution of the model presented in the previous section is (C.3). In order to take the model to the item-level data, we use the log first difference of an item's share ($\Delta \ln s_{igt}$) rather than its level. This change of variable, along with

fixed effects helps us deal with several econometric problems that may bias our estimates.[2] We will discuss these issues in detail below.

First, log-differencing (C.3) and substituting in (C.1), we arrive at

$$\Delta \ln \varphi_{igt} = \Delta \ln N_{gt} + (1-\sigma_g)\Delta \ln \left(\frac{p_{igt}}{P_{gt}}\right) + (\sigma_g - 1)\Delta\lambda_g(C_t) \ln \theta_{ig}. \tag{C.5}$$

To allow for estimation of (C.5), we need to take a stand on the functional form of $\lambda_g(C_t)$. As a baseline, we follow Hallak (2006), and assume that the quality parameter is linear in the log of total expenditures: $\lambda_g(C_t) = \eta_g + \mu_g \ln C_t$. We allow for heterogeneity in the average intensity of demand for quality of items in a group (η_g), as well as for the impact of income on quality demand across groups (μ_g). We then rewrite (C.5) as

$$\Delta \ln \varphi_{igt} = \Delta \ln N_{gt} + (1-\sigma_g)\Delta \ln \left(\frac{p_{igt}}{P_{gt}}\right) + (\sigma_g - 1)\mu_g \ln \theta_{ig}\Delta \ln C_t, \tag{C.6}$$

where the η_g disappears from taking first differences, and since the aggregate price index, P_t, is the numéraire, C_t is expressed in real terms.

Appendix D Identification and Additional Demand Estimation Results

D.1 Identification

The demand estimation presented in Section 4.1 faces several identification issues, which we discuss in this subsection, and address in further estimation results in the following subsection. First, rather than using the model-implied price index to derive group-level prices as a function of the item level prices, we compute the price indexes at the group level with the Törnqvist index. This approach may lead to measurement error due to unaccounted income-driven substitution, which would be picked up by the model-based group price index of (C.4). However, since $\Delta \ln P_{gt}$ enters the estimating equation linearly (both in the relative price and in deflating total expenditures), we eliminate this potential bias by including fixed effects that vary at the product group×time dimension.

Second, several papers have made the argument that trade costs went up during the crisis due to the freezing of trade credit, which made international trade more costly (Ahn, Amiti

[2]Note that by studying the growth rate of shares we are implicitly ignoring the impact of entry and exit on expenditure switching. We are not concerned with this omission given the importance of *intensive* margin – and correspondingly small role of the *extensive* margin – highlighted in Finding 1 of Section 3. Furthermore, we are able to control for changes in the number of items per product group each period by using appropriate fixed effects.

and Weinstein, 2011). Some firms (either domestic or foreign) may also have been driven out of business, thereby impacting the price level and supply of goods in a given product group. Furthermore, domestic and foreign goods within a given product group may also differ along other dimensions, such as durability or distribution (general availability) in the stores. All these differences between domestic and foreign goods may bias the estimation of (9). To control for these potential biases, we also consider a more demanding set of fixed effects, which are at the product group×origin×time level, where the origin is a dummy variable equal to one if the good is domestic, and zero if it is foreign.[3] The inclusion of these fixed effects will control for any unobserved heterogeneity of domestic and foreign items at the product group level. These effects will also control for potential shocks at a very disaggregated level in order to capture the general equilibrium impact of the shocks *within* a product group, and differential impacts of these shocks on domestic and foreign goods.

We also address the possibility that firms will discriminate their pricing depending on item-level characteristics. For example, firms may set prices higher for more desirable items (which are also more expensive), and pricing behavior may not respond to shocks symmetrically over time across different types of goods. Furthermore, other unobserved item-level characteristics (e.g., durability) might also bias the estimated price and income coefficients. To account for these potential biases, we augment (9) with item-level fixed effects in an additional regression specification.

Including both item-level and product group×time fixed effects deal with many potential omitted variables and unobserved shocks at very disaggregated levels. However, it is still possible that the unobserved demand shocks (ε_{igt}) are correlated with price changes over time at the item level. Furthermore, it is possible that there are other factors at work, which lower consumer's propensity to buy imports, regardless of relative price changes, which would bias our estimation results. For example, Latvians may have become more "patriotic" during the crisis, and thus slanted their consumption to domestic goods. We therefore adopt two instrumental variables strategies using subsets of the data as further checks.

The first approach exploits the variation in bilateral exchange rates for different foreign items as cost shocks. Though we are looking at retail prices, there is evidence that exchange rates pass-through to the consumer level and can serve as viable instruments at the retail level by providing a plausible source of exogenous price variation (e.g., see Campa and

[3] In other words, we replace α_{gt} with product group×origin×time fixed effects, α_{got}.

Goldberg, 2006; Goldberg and Hellerstein, 2008; Nakamura and Zeron, 2010). Although Latvia maintained a fixed exchange rate to the euro throughout the sample period, it had floating exchange rates with a significant share of importers.[4] We can therefore exploit variation in exchange rate movements across time and trading partners within a given product group with multiple foreign items. We also exploit heterogeneity in pass-through at the item level by interacting the exchange rate changes with an item's average unit value, since pass-through may vary for cheap and expensive goods. Therefore, the IV strategy exploits both cross-sectional and time-series variations in the data. Specifically, we instrument the item-level price changes with six lags of nominal exchange rate changes, and the interaction of these changes with the average unit value.[5]

The second approach follows Hausman (1996) and instruments Latvian goods' price changes with price changes of the same items in another market (so called "Hausman instruments"). In particular, we restrict our sample to a set of Latvian-produced goods that we also have price changes for in Estonia, Latvia's neighboring country. As long as *item-level* demand shocks are uncorrelated across countries, this strategy will help deal with potential biases – for example, it will deal with issues such as "patriotism" affecting Latvians' consumption patterns during the crisis, which would have led to increased consumption of Latvian goods, irrespective of relative price changes. It is important to note that the inclusion of product group×time effects pick up any common shocks hitting both Estonia and Latvia at more macro levels.

D.2 Additional Heterogeneous Coefficient Results

Besides our core results reported in Table 6 for the heterogeneous coefficient regressions, we also allow for the possibilities of non-linearities in the income effect and explore more stringent sets of fixed effects. Table A1 presents our core results for the weighted-mean coefficients from these regressions, where columns (1) and (2) replicate the main results in Table 6. Next, column (3) allows for the possibility of a non-linear income effect by including a squared term of the change in aggregate real consumption interacted with quality.

[4]The foreign sample includes imports from 40 countries, out of which 24 have floating exchange rates with Latvia. These 24 countries accounted for 40% of imports in our sample, with Poland, Russia and Sweden being the largest non-euro trading partners. For the median non-euro trading partner, the quarterly exchange rate viz. Lats fluctuated in the range of 34% on average over the sample period – the ranges for Poland, Russia and Sweden were 36%, 32% and 24%, respectively.

[5]The first-stage regression is $\Delta \ln(p_{igt}/P_{gt}) = \alpha_{gt} + \beta \ln \bar{p}_{ig} + \sum_{k=1}^{6} \delta_k \Delta NER^i_{t-k} + \sum_{k=1}^{6} \gamma_k \ln \bar{p}_{ig} \Delta NER^i_{t-k} + \epsilon_{igt}$, where ΔNER^i_{t-k} is the quarterly nominal exchange rate change of the Lat vs. the currency of country that item i is shipped from, lagged $t-k$ quarters, with $k = 1, \ldots, 6$. The cutoff of 6 quarters was chosen given further lags did not increase fit, nor were significant.

The coefficient for the non-linear term is insignificant, while the estimated price and income coefficients do not differ significantly from the baseline estimates in column (2). Therefore, it does not appear that non-linearities are a concern. We next control for the possibility that domestic suppliers reacted differently than foreign ones during the boom and ensuing crisis. For example, data show that producers in Latvia responded to the severe crisis by cutting production costs (e.g., wages), which may have lowered prices of domestic final goods, including food items, relative to their imported counterparts (Blanchard, Griffiths and Gruss, 2013; Kang and Shambaugh, 2013a). Furthermore, there may be unobserved time-varying differences between domestic and foreign goods within the narrowly defined product groups, which would bias our results, as discussed above. To investigate these possibilities, we re-estimate the NH model controlling for product group×origin×time fixed effects. These results are presented in column (4). The estimated relative price coefficients increase (in absolute value) relative to the coefficients in the baseline estimations of column (2), but the difference is marginal. The estimated income coefficient is smaller than that of the baseline estimation, but this difference is again marginal and statistically indistinguishable from zero. Column (5) next controls for item-level fixed effects in order to capture omitted time-invariant item-level characteristics. Controlling for these fixed effects increases the magnitude of the price coefficient, and decreases the income coefficient relative to the baseline estimation of column (2). This confirms the potential of estimation bias of the demand equation, but neither the price nor income coefficients vary dramatically in magnitude across columns (2) and (5).

D.3 Pooled Coefficient Results

We also explore restricting the price and income coefficients to being homogeneous (i.e., $\beta_{1g} = \beta_1 \ \forall g$, and $\beta_{2g} = \beta_2 \ \forall g$), in order to study whether results differ substantially from the heterogeneous coefficient estimates of Table 6 and Table A1. Table A2 presents our baseline estimations. Column (1) presents the CES model, where we only consider relative price changes and ignore potential income effects. The estimated coefficient is -2.390, and is significant at the 1% level. This coefficient implies a price elasticity, σ, equal to 3.390.[6] Column (2) presents the baseline results for the NH model. The estimated β_2 coefficient is positive and significant, with a value of 1.104, which implies a value of μ equal to 0.464. Columns (3)-(5) next present additional controls as in the heterogeneous coefficient results

[6]This elasticity is the same order of magnitude compared to previous estimates using retail level prices, such as for the coffee market (Nakamura and Zeron, 2010), or using scanner data across many goods (Handbury, 2013).

in Table A1 – the pooled estimates are similar to the specifications for our main regression results.

D.4 Instrumental Variables and Robustness Checks

Table A3 presents instrumental variable estimates for two sub-samples of data. Panel A uses data for the sample of imported items from non-euro countries, and exploits exchange rate variation with these trading partners.[7] The first two columns show results for the CES specification, while the latter two for the NH specification. Columns (1) and (3) run OLS regressions in order to compare with their IV counterparts in columns (2) and (4), respectively. Both sets of IV results have larger prices elasticities compared to the OLS ones (which are similar in magnitude to the estimates for the whole sample in Table A2), and are significant. Turning to the income elasticity in column (4), the IV estimate decreases slightly relative to the OLS one in column (3), but the difference is statistically indistinguishable from zero. Panel B next considers the subset of domestic goods that are also sold in Estonia, and uses the Estonian price change as the instrument.[8] This is a much smaller set of goods than the pooled sample, but we still have sufficient power to identify the impact of relative price changes and income on the within-group variation of expenditure shares. Again, we present OLS and IV estimates for the reduced sample for both the CES and NH models. Similar to the findings in Panel A, the CES and NH IV relative price coefficients are larger (in absolute value) than their OLS counterparts. Further, the income coefficient in column (4) is larger than the OLS one, while both coefficients are significant at the 10% level. Interestingly, the estimated coefficients are of the same order of magnitude relative to what we find in the other samples of data, and the income coefficients are not statistically different from each other.

In sum, the IV results across the two sub-samples broadly support the estimated income effects of the OLS regressions of Table A2, while the price elasticities are larger in absolute value. We explore the quantitative implications of the larger price elasticities in Section D.5, by comparing the predicted expenditure switching during the crisis using both the IV estimates and the baseline OLS coefficients estimated using the full sample of data.

[7] The long-run pass-through coefficient for the average unit value item, calculated as $\sum_{k=1}^{6}\left(\hat{\delta}_k + \ln \bar{p}\hat{\gamma}_k\right)$, is -0.05 (s.e. $= 0.022$) reflecting that a depreciation of the Lats leads to a price increase. Given that we are looking at retail prices, a small pass-through coefficient is not surprising (e.g., due to a large non-tradable component, see Burstein, Eichenbaum and Rebelo, 2005), but the null of weak instruments is rejected at conventional levels of significance, and the R-squared of the first-stage regression is 0.16.

[8] The first-stage coefficient for Estonian price changes is 0.13 (s.e. $= 0.013$), and the R-squared of the regression is 0.34. The null of weak instruments are rejected at conventional levels of significance.

Finally, Table A4 presents results for the baseline pooled NH regression model for a variety of sample splits, as well as investigating heterogeneity across different product group characteristics. In particular, we split the data by (i) domestic vs. foreign items (columns (1) and (2)); (ii) types of stores (columns (3)-(5));[9] and (iii) different non-parametric specifications to check for non-linear effects of relative price changes and income, where we interact quartiles defined by (a) product group foreign shares (column (6)); (b) variation of item-level unit values within a product group (column(7)); (c) average item-level unit values across product groups (column (8)), and (d) average difference in foreign and domestic unit values within a group (column (9)). In sum, all income and price coefficients are comparable to the baseline estimates throughout all the different cuts of the data.

D.5 Pooled Estimation Predictions

We provide a back-of-the-envelope calculation that applies the pooled estimation coefficients to predict aggregate expenditure switching during the crisis period. This exercise allows us to examine different bounds for the role of income-induced expenditure switching, given the IV estimates. For the back-of-the-envelope calculation, we draw on information from Section 3, along with moments from the data used in our regression analysis. In particular, according to Finding 2 the relative price change for imported goods was 0.004. The average relative unit value ($\ln \bar{p}_{ig}$) for foreign items was 0.25 over the crisis period, and the change in real consumption per capita ($\Delta \ln C_t$) was -0.13. Next, we use coefficients from our baseline estimate in column (3) of Table A2, and compare the predicted within-group expenditure switching with the one based on the most conservative coefficient estimates, which are the IV estimates from column (4) of Table A3.

The results for the quantification exercises are presented in Table A5, where Panel A presents the results based on the OLS coefficients, and Panel B's numbers are based on the IV estimates. The first column displays the coefficient estimates, while column (2) presents the predicted expenditure switching (and associated standard errors) for the (i) price effect, (ii) income effect, and (iii) total. Column (3) displays the share of expenditure switching that the price and income effects account for, along with their standard errors. The total predicted expenditure switching is very similar using both sets of estimates (-0.047 and -0.046 for the OLS and IV coefficients, respectively). Turning to the decompositions, the income effect explains roughly 78% of total expenditure switching based on the OLS

[9] This specification checks whether results are not simply being driven by consumers switching across stores as in Coibion, Gorodnichenko and Hong (2015).

coefficients. Given the potential estimation bias, this may be viewed as an upper bound. However, turning to the results based on the IV estimation, where the price elasticity is more than 1.5 times the size of the OLS estimate and the income coefficient is 20% smaller than it's OLS counterpart, one sees that the income effect still explains 63% of the predicted expenditure switching. Therefore, and as we shall see in the following section, the income effect plays an important role in driving expenditure switching within product groups.

Appendix E Predicted Aggregate Within Expenditure Switching and Standard Errors

The section outlines how we calculate the aggregated predicted expenditure switching, along with corresponding standard error bands. We first define the predicted value of the item share that we obtain from the regression (9). In particular, we are only interested in the predicted value due to either the change in prices or the change in income (quality effect) or both, so let $\beta_g \equiv [\beta_{1g}, \beta_{2g}]$, and $Z_{igt} \equiv [\Delta(p_{igt}/P_{gt}), \ln \bar{p}_{ig} \times \Delta C_t]'$, and ignore the group×time fixed effects.

Specifically, we predict the *aggregate* within-group y-on-y expenditure switching using coefficient estimates of the regression model (9), allowing for coefficient heterogeneity in the βs across product groups. We use the full distributions of estimated coefficients to calculate the predicted expenditure switching between any consecutive quarters τ and $\tau - 1$:

$$\widehat{(s^F_\tau - s^F_{\tau-1})}^{\text{Within}} = \sum_g s_{g\tau-1}(\widehat{\varphi}^F_{g\tau} - \varphi^F_{g\tau-1}), \tag{E.1}$$

where $\widehat{\varphi}^F_{g\tau}$ is generated using the following methodology:

1. Take the estimated coefficients from the within-group regressions (9), $\widehat{\beta}_{1g}$s and $\widehat{\beta}_{2g}$s, and predict the quarterly growth rate of every item i's share in group g sales $\Rightarrow$ $\widehat{\Delta \ln \varphi_{ig\tau}}$.

2. Use the quarterly growth rate to calculate the τ share of item i conditional on the item's share at $\tau - 1$ observed in the data, $\varphi_{ig\tau-1} \Rightarrow \widehat{\varphi}_{ig\tau}$.

3. Keep only foreign items' shares, and aggregate them within a group g to obtain the group-specific foreign share $\Rightarrow \widehat{\varphi}^F_{g\tau} = \sum_{i \in I^F_{g,\tau}} \widehat{\varphi}_{ig\tau}$.

The predicted q-on-q within-group expenditure switching is then cumulated into a y-on-y

measure by summing up four consecutive quarters in order to eliminate seasonality issues:

$$\widehat{(s_t^F - s_k^F)}^{\text{Within}} = \sum_{\tau=k}^{t} \sum_g s_{g\tau-1}(\widehat{\varphi}_{g\tau}^F - \varphi_{g\tau-1}^F), \tag{E.2}$$

where $k = t - 3$.

The following steps provide more details on the procedure, as well as how we calculate analytical standard errors for the predicted aggregate within-group expenditure switching between k and t. Note that we also construct a data counterpart, $(s_t^F - s_k^F)^{\text{Within}} = \sum_{\tau=k}^{t} \sum_g s_{g\tau-1}(\varphi_{g\tau}^F - \varphi_{g\tau-1}^F)$, to compare to the predicted values.[10]

E.1 Step 1

We use the estimated coefficient to predict the growth rates of item shares at any quarter τ:

$$\widehat{\Delta \ln \varphi}_{ig\tau} = \widehat{\beta}_g Z_{ig\tau}, \tag{E.3}$$

where $\widehat{\beta}_g \sim \text{N}\{\beta_g, \Sigma_g\}$, and we have estimates of Σ_g, $\widehat{\Sigma}_g$, which are based on clustering.

E.2 Step 2

Next, we take actual data at time $\tau - 1$ and use (E.3) to predict the within-group share of item i at any quarter τ:

$$\widehat{\varphi}_{ig\tau} = \left(1 + \widehat{\beta}_g Z_{ig\tau}\right) \varphi_{ig\tau-1}. \tag{E.4}$$

Note that the randomness of $\widehat{\varphi}_{ig\tau}$ comes from $\widehat{\beta}_g$.

E.3 Step 3

We next simply aggregate (E.4) for each group g for only foreign items to obtain a product group's predicted foreign share:

$$\begin{aligned} \widehat{\varphi}_{g\tau}^F &= \sum_{i \in I_{g\tau/\tau-1}^F} \left(1 + \widehat{\beta}_g Z_{ig\tau}\right) \varphi_{ig\tau-1} \\ &= \sum_{i \in I_{g\tau/\tau-1}^F} \varphi_{ig\tau-1} + \sum_{i \in I_{g\tau/\tau-1}^F} \left(\widehat{\beta}_g Z_{ig\tau}\right) \varphi_{ig\tau-1} \\ &= \varphi_{g\tau-1}^F + \sum_{i \in I_{g\tau/\tau-1}^F} \left(\widehat{\beta}_g Z_{ig\tau}\right) \varphi_{ig\tau-1}. \end{aligned} \tag{E.5}$$

Note here that the foreign share for a given group is going to depend on the previous period's observed foreign share. Further, for the next step, define $Q_{g\tau}^F \equiv \sum_{i \in I_{g\tau/\tau-1}^F} Z_{ig\tau} \varphi_{ig\tau-1}$.

[10] This measure is not identical to the within expenditure switching presented in Figure 5, but the two series are very similar.

E.4 Step 4

Calculate the model predicted expenditure switching between periods τ and $\tau-1$:

$$\widehat{(s_{\tau}^{F}-s_{\tau-1}^{F})}^{\text{Within}}=\sum_{g}s_{g\tau-1}(\widehat{\varphi}_{g\tau}^{F}-\varphi_{g\tau-1}^{F}), \tag{E.6}$$

which we then aggregate over four consecutive quarters to arrive at predicted year-on-year expenditure switching:

$$\widehat{(s_{t}^{F}-s_{k}^{F})}^{\text{Within}}=\sum_{\tau=k}^{t}\sum_{g}s_{g\tau-1}(\widehat{\varphi}_{g\tau}^{F}-\varphi_{g\tau-1}^{F}), \tag{E.7}$$

where $k=t-3$.

E.5 Aggregate Variance

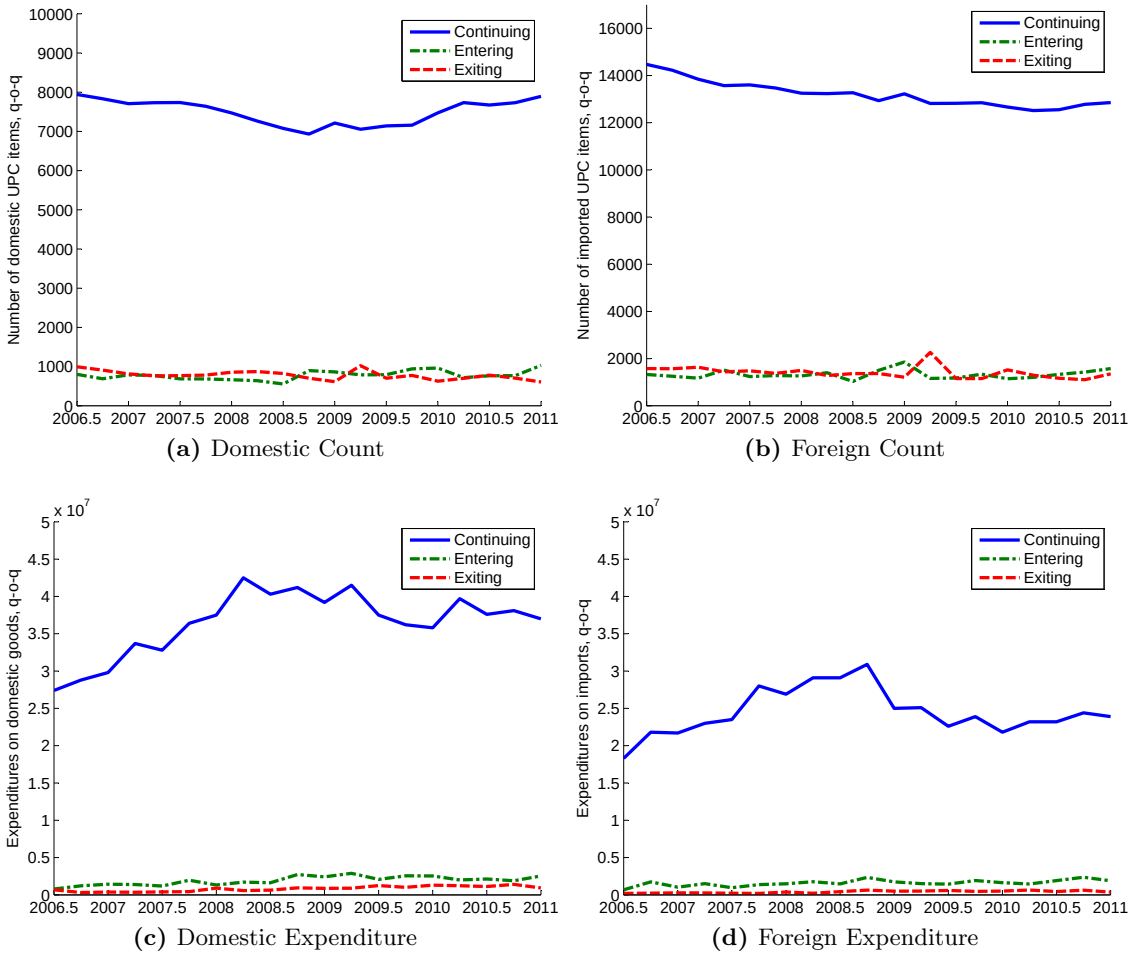
We are interested in calculate the variance of (E.7):

$$\text{Var}\left\{\sum_{\tau=k}^{t}\sum_{g}s_{g\tau-1}\widehat{\varphi}_{g\tau}^{F}\right\}=\sum_{\tau=k}^{t}\text{Var}\left\{X_{\tau}\right\}+2\sum_{p\neq\tau}^{t}\sum_{\tau=k}^{t}\text{Cov}\left\{X_{\tau},X_{p}\right\}, \tag{E.8}$$

where $X_{\tau}=\sum_{g}s_{g\tau-1}\widehat{\varphi}_{g\tau}^{F}$, and $X_{p}=\sum_{g}s_{gp-1}\widehat{\varphi}_{gp}^{F}$, or

$$\begin{aligned}\text{Var}\left\{\sum_{\tau=k}^{t}\sum_{g}s_{g\tau-1}\widehat{\varphi}_{g\tau}^{F}\right\}&=\sum_{\tau=k}^{t}\sum_{n}\sum_{g}s_{g\tau-1}^{2}\left(Q_{g\tau}^{F}\right)^{2}\text{Cov}\left\{\widehat{\beta}_{g},\widehat{\beta}_{n}\right\}\\&\quad+2\sum_{p\neq\tau}\sum_{\tau=k}^{t}\sum_{n}\sum_{g}s_{g\tau-1}s_{gp-1}Q_{g\tau}^{F}Q_{np}^{F}\text{Cov}\left\{\widehat{\beta}_{g},\widehat{\beta}_{n}\right\}\\&=\sum_{p=k}^{t}\sum_{\tau=k}^{t}\sum_{n}\sum_{g}s_{g\tau-1}s_{gp-1}Q_{g\tau}^{F}Q_{np}^{F}\text{Cov}\left\{\widehat{\beta}_{g},\widehat{\beta}_{n}\right\}\end{aligned} \tag{E.9}$$

Figure A1. Domestic and Import Goods: Continuing, Entry, and Exit

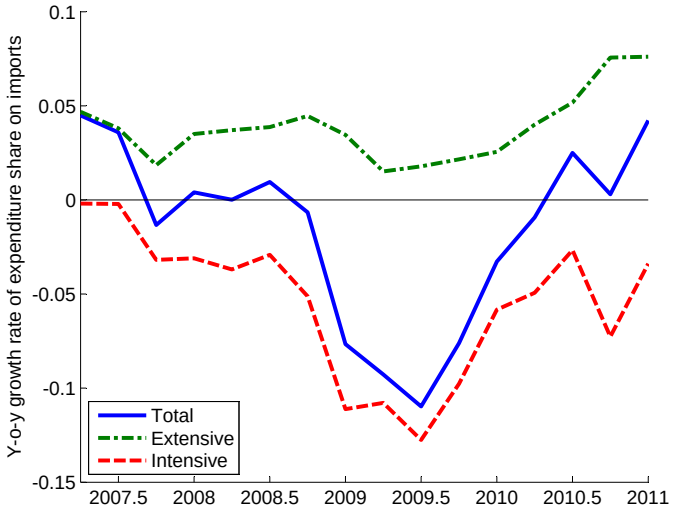


(a) Domestic Count **(b)** Foreign Count

(c) Domestic Expenditure **(d)** Foreign Expenditure

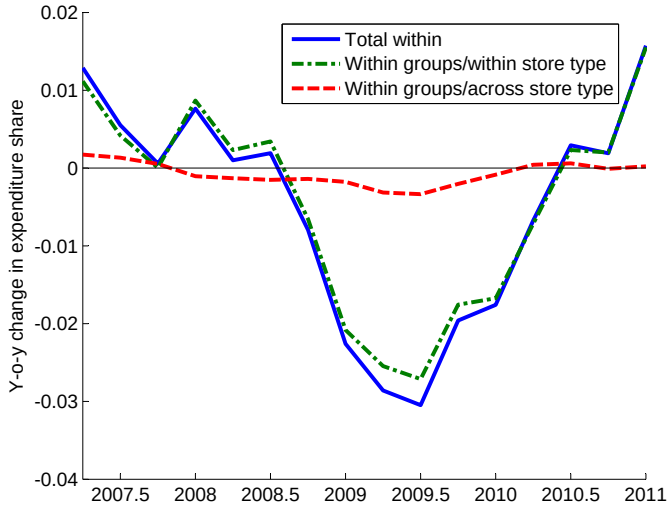
Notes: This figure plots the time series of items that (i) continue, (ii) enter, and (iii) exit from one quarter to the next for domestic and foreign goods. The top two panels present the count of UPCs, while the bottom two panels present total expenditures on the types of goods.

Figure A2. Growth Rate of Expenditure Switching: Total and Intensive and Extensive Margins



Notes: This figure plots the growth rate of the total expenditure share on imported goods, as well as the contribution to growth due to changes for continuing goods – the intensive margin – and due to net entry and exit of goods – the extensive margin. Growth rates are calculated using quarterly data and are then accumulated over a four-quarter overlapping rolling window.

Figure A3. Decomposition of Within Expenditure Switching: Within Store Type and Across Store Type Components



Notes: This figure decomposes total expenditure switching within 4-digit product groups into two sub-components: (i) expenditure switching within product groups and within store types, and (ii) expenditure switching within product groups but across store types. Changes in expenditure shares are expressed as y-o-y changes, based on quarterly data.

Table A1. CES and Non-Homothetic Models' Heterogeneous Coefficient Regressions: Weighted-Mean Estimates

	(1)	(2)	(3)	(4)	(5)
$\Delta \ln(p_{igt}/P_{gt})$	-2.938	-2.949	-2.956	-2.971	-3.032
	(0.028)	(0.027)	(0.027)	(0.028)	(0.027)
$\ln \bar{p}_{ig} \times \Delta \ln C_t$		1.701	1.531	1.740	1.581
		(0.140)	(0.155)	(0.154)	(0.143)
$\ln \bar{p}_{ig} \times (\Delta \ln C_t)^2$			-3.576		
			(2.366)		
Observations	372,484	372,484	372,484	372,484	372,484
Group×time F.E.	7,344	7,344	7,344	-	7,344
Group×origin×time F.E.	-	-	-	11,638	-
Item F.E.	-	-	-	-	26,555
R^2	0.143	0.146	0.145	0.179	0.319

Notes: This table presents weighted means of the coefficients of regression model (9), and is an extension of the heterogeneous coefficient regressions reported in the main regression table Table 6, where columns (1) and (2) are identical to the results in the main text. The weights are a product group's share of total expenditures over the sample period. Column (1) presents the price coefficients for the CES model; columns (2) and (3) present the price and income coefficients, for the non-homothetic model with and without a quadratic term for income, respectively. These specifications are run with product group×time fixed effects. Columns (4) and (5) run the baseline NH model with product group×origin×time or item-level fixed effects, respectively. Standard errors clustered at the item level are in parentheses.

Table A2. CES and Non-Homothetic Pooled Regression Estimates: Baseline

	(1)	(2)	(3)	(4)	(5)
$\Delta \ln(p_{igt}/P_{gt})$	-2.390	-2.391	-2.391	-2.413	-2.530
	(0.020)	(0.020)	(0.020)	(0.020)	(0.020)
$\ln \bar{p}_{ig} \times \Delta \ln C_t$		1.104	1.122	1.030	0.973
		(0.066)	(0.075)	(0.070)	(0.067)
$\ln \bar{p}_{ig} \times (\Delta \ln C_t)^2$			0.540		
			(1.041)		
Observations	372,484	372,484	372,484	372,484	372,484
Group×time F.E.	7,344	7,344	7,344	-	7,344
Group×origin×time F.E.	-	-	-	11,638	-
Item F.E.	-	-	-	-	26,555
R^2	0.120	0.120	0.120	0.154	0.295

Notes: This table presents coefficients of the pooled estimation of regression model (9), and correspond to the heterogeneous coefficient regressions reported in the main regression table Table 6, and Table A1. Column (1) presents the price coefficients for the CES model; columns (2) and (3) present the price and income coefficients, for the non-homothetic model with and without a quadratic term for income, respectively. These specifications are run with product group×time fixed effects. Columns (4) and (5) run the baseline CES and NH models with product group×origin×time fixed. Standard errors clustered at the item level are in parentheses.

Table A3. CES and Non-Homothetic Pooled Regression Estimates: Instrumental Variables

Panel A: Foreign Items

	CES		*NH*	
	(1)	(2)	(3)	(4)
	OLS	IV	OLS	IV
$\Delta \ln(p_{igt}/P_{gt})$	-2.410	-4.177	-2.410	-3.949
	(0.045)	(0.628)	(0.045)	(0.621)
$\ln \bar{p}_{ig} \times \Delta \ln C_t$			0.875	0.873
			(0.141)	(0.145)
Observations	90,323	90,323	90,323	90,323
Group×time F.E.	4,468	4,468	4,468	4,468
R^2	0.198	0.158	0.198	0.168
Instrument	-	NER & Quality	-	NER & Quality
Kleibergen-Paap F-stat	-	13.18	-	13.20

Panel B: Subset of Domestic Items

	CES		*NH*	
	(1)	(2)	(3)	(4)
	OLS	IV	OLS	IV
$\Delta \ln(p_{igt}/P_{gt})$	-2.547	-2.712	-2.551	-2.741
	(0.074)	(0.399)	(0.074)	(0.402)
$\ln \bar{p}_{ig} \times \Delta \ln C_t$			0.867	0.921
			(0.416)	(0.431)
Observations	19,119	19,119	19,119	19,119
Group×time F.E.	2,060	2,060	2,060	2,060
R^2	0.458	0.457	0.458	0.457
Instruments	-	Estonian prices	-	Estonian prices
Kleibergen-Paap F-stat	-	104.1	-	102.8

Notes: This table presents coefficients of the pooled estimation of regression model (9), instrumenting for the change in the relative price. Panel A uses data for imported items from non-euro countries, and instruments using six lags of exchange rate changes, and these changes interacted with an item's average unit value (relative to the group price index). Panel B uses a subset of domestic items that are also sold in Estonia, and instruments using contemporaneous values of Estonian items' price changes. These specifications are run with product group×time fixed effects. Standard errors clustered at the item level are in parentheses.

Table A4. Pooled Regression Estimates: Robustness

	(1) Foreign	(2) Domestic	(3) D Store	(4) S Store	(5) H Store	(6) F share	(7) St.Dev(UV)	(8) Mean(UV)	(9) UV_F-UV_H
$\Delta \ln(p_{igt}/P_{gt})$	-2.522	-2.633	-1.391	-2.490	-2.703				
	(0.026)	(0.032)	(0.102)	(0.032)	(0.026)				
$\ln \bar{p}_{ig} \times \Delta \ln C_t$	0.862	0.984	1.052	1.062	0.821				
	(0.082)	(0.136)	(0.362)	(0.125)	(0.079)				
Q1 $\Delta \ln(p_{igt}/P_{gt})$						-2.729	-2.678	-2.230	-2.552
						(0.040)	(0.039)	(0.047)	(0.041)
Q2 $\Delta \ln(p_{igt}/P_{gt})$						-2.982	-2.617	-2.792	-2.097
						(0.042)	(0.036)	(0.038)	(0.037)
Q3 $\Delta \ln(p_{igt}/P_{gt})$						-2.678	-2.744	-2.767	-2.896
						(0.038)	(0.042)	(0.039)	(0.044)
Q4 $\Delta \ln(p_{igt}/P_{gt})$						-2.081	-2.139	-2.230	-3.107
						(0.053)	(0.043)	(0.040)	(0.050)
Q1 $\ln \bar{p}_{ig} \times \Delta \ln C_t$						0.454	1.846	0.384	1.013
						(0.153)	(0.209)	(0.181)	(0.175)
Q2 $\ln \bar{p}_{ig} \times \Delta \ln C_t$						1.316	0.947	3.197	0.440
						(0.139)	(0.171)	(0.816)	(0.143)
Q3 $\ln \bar{p}_{ig} \times \Delta \ln C_t$						1.065	1.122	1.708	1.462
						(0.111)	(0.126)	(0.259)	(0.145)
Q4 $\ln \bar{p}_{ig} \times \Delta \ln C_t$						1.090	0.766	1.158	1.132
						(0.216)	(0.097)	(0.090)	(0.121)
Observations	237,036	135,155	21,802	133,184	212,031	322,774	372,484	372,484	322,774
Group×time F.E.	6,258	5,379	3,887	6,774	7,207	5,623	7,344	7,344	5,623
Item F.E.	17,824	8,382	3,294	15,010	23,756	23,061	26,555	26,555	23,061
R^2	0.312	0.349	0.488	0.345	0.326	0.304	0.297	0.297	0.305

Notes: This table presents coefficients of the pooled estimation of regression model (9) for a variety of sample splits and robustness checks for the NH model. Columns (1) and (2) present the price and income coefficients splitting the sample into Foreign and Domestic items only, respectively. Columns (3)-(5) splits the sample across the three different store types, where the 'D Store' is the discount market, the 'M Store' is a typical supermarket, and the 'H Store' is the hypermarket. Columns (6)-(9) include interactions of quartile dummies (Q_1-Q_4) with the relative price and income variables. Column (6) defines the quartiles by the average foreign share of expenditures of a product group over total expenditures during the sample period. Column (7) defines the quartiles by the standard deviation of goods' unit values within a product group. Column (8) defines the quartiles by the mean of goods' unit values (relative to its product groups unit value) over the sample period. Column (9) defines quartiles by the average difference of foreign and domestic unit values within a product group over the sample period. These specifications are run with product group×time fixed effects. Standard errors clustered at the item level are in parentheses.

Table A5. Within-Group Predicted Expenditure Switching for OLS and IV Pooled Estimates

Panel A: Baseline OLS

	(1) Coef.	(2) Predicted ES	(3) Share
Price	-2.391	-0.010	0.221
	(0.020)	(0.000)	(0.010)
Income	1.104	-0.037	0.779
	(0.066)	(0.002)	(0.010)
Total	-	-0.047	
		(0.002)	

Panel B: Exchange Rate-Quality IV

	(1) Coef.	(2) Predicted ES	(3) Share
Price	-3.949	-0.017	0.371
	(0.621)	(0.003)	(0.054)
Income	0.873	-0.029	0.629
	(0.145)	(0.005)	(0.054)
Total	-	-0.046	
		(0.006)	

Notes: This table presents predicted within-group expenditure switching ('Predicted ES') for OLS and IV coefficient estimates. Panel A uses the baseline OLS coefficients from column (3) of Table A2, while Panel B uses the IV estimates based on the foreign subsample from column (4) of Table A3. Column (1) contains the estimated coefficients, column (2) shows the predicted expenditure switching, and column (3) presents the share of expenditure switching due to the price and income effects. The following values, corresponding to the crisis period and foreign items, were used to predict expenditure switching: $\Delta \ln(p_{igt}/P_{gt}) = 0.004$; $\Delta \ln C_t = -0.133$, and $\ln \bar{p}_{ig} = 0.25$. Standard errors clustered at the item level are in parentheses.

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