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INCOME-INDUCED EXPENDITURE SWITCHING

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ABSTRACT

Income-Induced Expenditure Switching*

This paper shows that an income effect can drive expenditure switching between domestic and foreign goods. We use a unique Latvian scanner-level dataset for food and beverages, covering the 2008-09 financial crisis, to study (i) relative price movements, and (ii) expenditure switching between domestic and imported goods. We document several empirical findings. First, imports contracted by 26%, with expenditure switching accounting for one third of the fall, while the relative price of foreign goods viz. the food CPI increased by 4.4% during the crisis. Second, the majority of the switching took place between items within narrowly defined product groups, while the relative price adjustment was across product groups. This puzzling asymmetry in expenditure and price adjustments, combined with a finding that consumers substituted towards lower unit value domestic items during the crisis, motivate us to model non-homothetic consumer demand. Over the crisis period, the estimated model explains eighty percent of the observed expenditure switching, which was driven almost entirely by an income effect.

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1 Introduction

Conventional international macroeconomics theory posits that changes in the relative price between domestic and foreign goods is the sole driver of expenditure switching. For example, following a currency devaluation, a country's imports fall and its exports rise in proportion to its relative price change vis-à-vis that of the rest of the world. This paper revisits external adjustment and relative price changes during a sudden stop episode empirically, and shows that an *income effect* can also play an important role in expenditure switching.

We examine the 2008–09 sudden stop crisis in Latvia, during which the country had limited nominal exchange rate flexibility, using a unique goods-level dataset. First, we quantify how much expenditure switching took place between domestic and imported goods. Second, we exploit these detailed data to examine at what margins did expenditure switching and relative price movements take place. Finally, we ask whether the observed relative price changes explained the observed expenditures switching through the lens of 'standard' economic models, or are other channels required in order to match the data.

We are able to measure relative price and consumption changes across goods by using a scanner-level dataset on food and beverages, which covers twenty quarters over the 2006Q2–2011Q1 period, where Latvia experienced a boom-bust episode.[1] These data provide both prices and quantities at the individual item level, and crucially identify the country origin of each good, and detailed product groups to which items belong to. Using the data, we find that during the crisis period real consumption of imports fell by 26%. We then use the item-level dimension of the data to present three main findings, which help explain this fall, and motivate the modeling and estimation strategy in the remainder of the paper.

First, we find that expenditure switching from imported to domestic goods accounted for one-third of the total fall in imports observed in the dataset.[2] The majority of this expenditure switching was driven by substitution between goods *within* narrowly defined product groups. Second, this expenditure switching was accompanied by a 4.4% rise in the price of imports relative to the food CPI, where this change was almost entirely driven by changes in prices *across* product groups. Third, within product groups, consumers substituted towards cheaper domestic items during the crisis, irrespective of any relative price change.

The asymmetry in adjustment of expenditures and relative prices within and across

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[1] See Blanchard et al. (2013) for a forensic account of Latvia's boom, bust, and recovery over 2000–13.

[2] The other two-thirds is due to a proportional fall in domestic and imported goods in response to the crisis-induced fall in aggregate income.

product groups (Findings 1 and 2), along with the average domestic/foreign unit value differential and substitution between domestic and foreign items (Finding 3), motivate us to set up a demand-side model of the economy, where a household's consumption decision not only depends on the good's relative price, but also on other item characteristics. We follow the framework in Hallak (2006), where a non-homothetic term enters the utility function via an item's quality, such that the consumer's intensity for demand of quality varies with her income level. Though we do not explicitly model quality, we follow the literature and allow for the unit value of an item to proxy for this "quality" term. What is crucial is that we want to capture the possibility for a consumer to substitute between comparable cheap and expensive items, irrespective of their relative price change, when hit by an income shock, such as the one consumers faced during Latvia's crisis.[3]

The model provides us with a simple structure, which we apply in a panel regression setting to estimate the impact of relative price and income changes on items' expenditure shares. Our empirical strategy exploits the disaggregated item-level data, and allows us to control for a host of potential factors that might otherwise bias the estimation results. We then use the estimated parameters to construct a predicted aggregate measure of total expenditure switching over the sample period. The results are quite striking. First, the baseline CES model, which only allows for a relative price change channel, and which underlies the workhorse models in international macroeconomics, performs poorly: though the estimated relative price parameters are significant, the model's predicted expenditure switching does not match the switching observed in the data, particularly during the crisis episode. Second, the non-homothetic model, which nests the baseline CES model, is better able to match observed expenditure switching during the crisis – it captures eighty percent of what is observed in the data. Therefore, allowing for an income channel, where consumers can substitute towards cheaper goods when their income falls steeply, helps explain Latvia's expenditure switching on the import side during the crisis.

Note that neither the theoretical model nor the estimation differentiate between domestic or foreign items. Therefore, what is the intuition for the aggregate results on expenditure switching and the role of relative prices and income? Finding 3 points to one possible answer: foreign items are on average more expensive than domestic ones in Latvia. Therefore, given the non-homothetic channel, when Latvian consumers substituted to cheaper items

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[3]Recent empirical work in the international trade literature shows that rich/poor countries tend to be net exporters/importers of higher quality (unit value) items (Schott, 2004; Hummels and Klenow, 2005; Choi et al., 2009; Hallak and Schott, 2011; Feenstra and Romalis, 2013). These results have motivated theoretical work, which endogenizes the non-homotheticity of a country's demand for quality (see Fajgelbaum et al., 2011, for example).

during the crisis, they also moved away from foreign ones. This finding is consistent with the "flight from quality" hypothesis put forth in Burstein et al. (2005), who present facts on consumer shopping patterns during the 2001 Argentinean crisis. We are able to shed further light on the impact of this "flight" by identifying a movement from foreign to domestic items in our dataset. Furthermore, Latvia did not experience a large exchange rate change, which allows us to more cleanly identify the potential for income-induced switching.

The literature to which this paper most directly contributes to has focused on relative price adjustment as the driver of the external adjustment after sudden stops. Our contributions are twofold. First, we provide evidence for relative price adjustment under a pegged regime. There is a large literature examining international prices and exchange rates (see Burstein and Gopinath, 2013, for a recent survey), but only limited evidence has been provided examining price movements under a fixed exchange rate (e.g., see Parsley and Popper, 2006, for the case of Hong Kong), though recent work has contrasted real exchange rate adjustments in and outside the Eurozone (Berka et al., 2012). Second, we measure expenditure switching and assess its drivers. Though the *theoretical* literature on expenditure switching and the role of the exchange rate policy is large (see Engel (2003) for a review; and Burstein et al. (2005), Kehoe and Ruhl (2008), Mendoza (2005), Obstfeld and Rogoff (2005) for work studying sudden stop episodes), the *empirical* literature on measuring the impact of relative price changes on expenditure switching using both domestic and foreign quantity and price data at the microeconomic level is, to the best of our knowledge, non-existent.[4]

Diaz Alejandro (1965) is an early study of how income effects can affect external reblancing.[5] He investigates how consumption behavior differences across the income distribution within the economy – specifically the difference between wage and non-wage earners – can affect the demand of different sectors' imports. Unfortunately, we do not have household data to match to our scanner data in order to examine how Engel's Law played a role in Latvia's external adjustment. Finally, a potential "flight from quality," which would show up as income driven expenditure switching for a small emerging market, like Latvia, may have been a more general phenomenon across countries during crises – some recent work has pointed to a fall in the quality composition of large EU countries' exports during the financial crisis (Berthou and Emlinger, 2010; Esposito and Vicarelli, 2011).

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[4]Of course, there is a long-standing literature that estimates import elasticities, which has more recently highlighted the importance of heterogeneity across sectors. See, Imbs and Méjean (2009) and Feenstra et al. (2012) for two recent contributions. There is also the literature studying the possibility of a "J-Curve" in the trade balance following an exchange rate change, which follows the classic work of Magee (1973).

[5]We thank Chang-Tai Hsieh for bringing Diaz Alejandro's work to our attention.

The rest of the paper is organized as follows. Section 2 describes the data. Section 3 presents empirical findings, which are used to motivate the modeling and empirical analysis of the Latvian experience. Section 4 presents the model. Section 5 estimates the model, and quantifies the contribution of relative price changes and income effects in explaining the observed expenditure switching. Section 6 concludes.

2 Data Description

The analysis is based on detailed scanner-level data, which contain monthly information on quantities sold and the average price level charged for 13-digit UPC items sold by one of Latvia's largest retailers. The data are collected across three types of stores that the retailer owns and runs: (i) a 'Hypermarket' (HM) (ii) a 'Supermarket' (SM) and a 'Discounter' (D).[6] Each store type's data are aggregated across the respective type's sales-per-item across the country, so there is no geographical distinction by type of store. In total, there are over 100,000 UPC-store pair items, and 64,000 unique items, covering the six-year period May 2006–May 2011. The coverage of goods is primarily for food and beverages (F&B), but the dataset also contains other consumer non-durables, such as toiletries. Besides quantity and price information, the dataset also provides information on the type of unit and the net content of each UPC item.

The retailer provides 2-, 3-, and 4-digit classifications of the items into product groups. An example of a 2-digit product group would be 'hot drinks,' which at the 3-digit level is further broken down into 'tea,' 'coffee,' and 'cacao.' The 3-digit group 'tea' is further broken down at the 4-digit level into types of tea. For example, there is 'unflavored black tea,' 'flavored black tea,' 'herbal tea,' 'fruit tea,' etc.

The retailer also provides an item description with an accompanying retailer assigned 'material' code. An example would be "SOY SAUCE BLUE DRAGON LIGHT 150ML; 'material' code: 111455." A given 'material' code can be assigned to multiple UPC items. This would be the case, for example, if a good's label is updated, but there is no change in the item's name or net content.

The UPC is crucial for the analysis because it allows us to identify the domestic/foreign origin of each item. In particular, the first three digits of the bar code identify the country in which the label was applied for. Because Latvia is a small market, foreign suppliers usually do not relabel their goods in Latvian. Instead, imported items carry a source country label

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[6]The SM and HM carry a wider variety of goods than the D, and the same good can vary in price across the three types of stores.

or a label intended for a larger destination market. This allows us to use the item's label to identify domestic/foreign origin. However, for items of foreign origin the label does not necessarily identify the country of production.[7]

An alternative approach to identifying the origin suggests that the UPC is a valid proxy. We zoom in on domestic/imported origin for a subset of 4-digit product groups that explicitly group items by origin (e.g., imported and domestic beer). Such product groups account for 11.6% of total F&B expenditures in our sample, 6.2% of which are identified as local and 5.4% as foreign. We find that for product groups that are identified as local, 97.3% of expenditures carry local UPCs. For product groups that are identified as imported, 97.2% of expenditures carry foreign UPCs. This suggests that for a small market, such as Latvia, UPCs can correctly identify the origin for more than 97% of expenditures.

One set of items for which we cannot identify the origin for all items is 'store products.' Such items are produced/labeled by the retailer, with the bulk of the goods falling into product groups such as 'store bake,' 'fruits and berries,' 'vegetables and root crops,' and 'fresh/processed meat and fish.' The UPC identifies such items as 'store products,' but provides no information about the origin of ingredients. Store items account for 18% of total food expenditures in the data. Over time there is a gradual decline in the weight of such items, from 20% in 2006 to 16% in 2011, but we find no evidence that this weight is affected by the crisis. We are able to identify store products at the 4-digit level when the group label explicitly contain either "imported" or "domestic." We are forced to drop all other (the bulk of) store products.

We also exclude items with Estonian and Lithuanian product labels and product groups dominated by such items, because the two economies went through a crisis very similar in magnitude to Latvia's. For the purpose of this paper one might expect items from these neighboring economies to behave like domestic rather than imported products. These two countries together account for 6% of expenditures with no significant trend over time.

Altogether these exclusions reduce the scanner data expenditures by 34%, out of which 24% of the reduction is due to dropping store products and goods from Estonia and Lithuania, and the remaining 10% is due to dropping whole product groups dominated by such items. This leave us with 37 2-digit product groups.

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[7]For example, the UPC of a bottle of tequila produced in Mexico, but labeled in the United States, and then shipped to Latvia would identify the bottle as originating from the United States.

2.1 Data Cleaning and Consolidation

As with any large micro dataset, data cleaning is needed. First, we drop items without a UPC. Second, we drop items where either quantity or price is less or equal to zero. Third, we drop items with the 0.05% largest price changes. Imposing these three conditions left total revenue virtually unchanged, decreasing it by 0.3%.

We next consolidate scanner data for homogeneous items, which improves the measurement of items' prices and entry/exit rates. We start by consolidating data by the triplet of (i) UPC (i), (ii) store type (s), and (iii) time period (t), because information pertaining to a given triplet can be reported in multiple entries. The consolidation is done by summing quantities, q_{ist}, and expenditures, x_{ist}, over identical triplets and then re-computing the unit values from aggregated data. As a check that the data we consolidate pertain to homogeneous items, we compare prices for all identical triplets and find that in 99.7% of cases prices are indeed identical.

On some occasions the UPC is an "overly" unique identifier of homogeneous items. For example, this would be the case if an item's label is frequently updated. Two such cases are presented in the panel below, which shows data entries as they appear in the dataset before aggregation:

4-digit product code	Product group description	'Material' code	Item description	Net content	UPC	Quantities 2009/4	Quantities 2009/5	Quantities 2009/6	Average prices 2009/4	Average prices 2009/5	Average prices 2009/6
6439	Other dental care eq	404199	DENTAL FLOSS ORAL B SATIN 25M	25 M	5010622017947	93	105	106	2.04	2.04	2.04
6439	Other dental care eq	404199	DENTAL FLOSS ORAL B SATIN 25M	25 M	5010622018258	20	5	9	2.04	2.04	2.04
2101	Fat-free milk	211961	MILK VALMIERA 0,5% 1L	1 L	4750074000500	1331	640		0.47	0.47	
2101	Fat-free milk	211961	MILK VALMIERA 0,5% 1L	1 L	4750074005062		994	2152		0.47	0.47

Items identified by the retailer's 'material' codes 404199 and 211961 have identical (i) product description, (ii) net content, (iii) average monthly prices, and (iv) producer code (identified by the first 6 digits of the UPC), but have varying 13-digit UPCs. For the purpose of this paper such items can be treated as homogeneous.

Motivated by this example, we consolidate data by the pair of (i) 'material' code and (ii) store type, when prices are identical in all periods for overlapping pairs. This consolidation decreases the number of unique UPCs in our sample by 12%.[8]

Lastly, we examine item homogeneity across the three store types. We find that for SMs and HMs, 70% of overlapping monthly prices of identical UPCs are identical, i.e., $p_{i,SM,t} = p_{i,HM,t}$, and in 97% of cases the deviation is less than 5%. The mean of this

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[8]This aggregation across items can be used to shed further light on the quality of the scanner dataset. If the UPC identifies unique items, then the multiple entries for the same UPC should all be assigned to the same 4-digit product group and have the same net content. We find that this is the case for 98.8–99.5% of aggregated items, depending on the store type.

price differential is 0.0007, and the median of the distribution is 0. Thus, there is strong evidence that items with identical UPCs in these two types of stores are homogeneous for our purpose. We therefore aggregate these UPC-store item pairs into a common UPC item. This consolidation does not change the number of unique UPCs in our sample, but reduces the store types to two – market (M) and discounter (D) – and decreases the number of unique UPC-store pairs by 29% percent. Price levels in Ds differ from the Ms. The mean of this price differential is –0.13, while the median is –0.11. Therefore, in this case, we continue to treat identical UPCs as different items, depending on whether the item is sold in the aggregated M stores or D stores.

2.2 Summary Statistics

Table 1 presents annual data on total sales for all products, as well as domestic and foreign goods separately. Given the sample period, we drop the last month of the sample, and define a year as May to April. So, for example, 2006 would be the year covering May 06–April 07. Looking at Columns (1) and (2), one sees that the value of sales increased until 2008–09 when the crisis hit, and there is then a pick up in 2010–11 as the Latvian economy began to recover. The same pattern holds for both domestic and foreign sales.

Next, Table 2 presents summary statistics for 2-digit product groups over the whole sample period. The 'Share' column reports the share of each product groups sales viz. total sales over the period, while the 'Foreign Share' column measures the foreign content of a given product group. There is considerable heterogeneity in both the size and foreign content of product groups at this relatively high level of aggregation. 'Alcoholic products' make up the largest value of total sales, accounting for 14.97% of aggregate revenue. Though the food and beverage sector is generally considered a tradable sector (Berka and Devereux, 2011; Crucini et al., 2005), we find considerable heterogeneity in import intensity among product groups. Foreign content ranges from as lows such as 0 ('Eggs') to a high of 0.99 ('Baby food'). While foreign share of total food sales is 37%, imports account for mere 2% of the least import-intensive product groups that make up a quarter of all food sales.

2.3 Aggregating Data into Macroeconomic Indicators

Food and beverages account for approximately 30% of total household expenditures in Latvia,[9] therefore the scanner data cover an important component of total consumption.

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[9]According to the Latvian CPI calculations, food has a 35% weight, but in the national income accounts data, F&B account for 25% of household expenditures. We therefore take a simple average to arrive at the 30%.

Furthermore, given the size of the retailer, the scanner dataset directy adds up to 15% of *aggregate* household expenditure on F&B over the period. In order to draw aggregate implications from the dataset, we next compare key aggregate statistics on F&B with equivalent series constructed from the scanner data.

First, as Figure 1 shows, the constructed aggregate price index closely mimics F&B's CPI.[10] Second, retail market share data, kindly provided to us by IGD Retail Analysis, show that during 2007–11 the retailer maintained a stable grocery retail market share of around 20%:

<div align="center">

Grocery Retail Market Share

Year	2007	2008	2009	2010	2011
Market share, %	20.0	21.7	22.5	22.4	21.5

</div>

Finally, Figure 2 plots the total revenue of foreign products across all stores and aggregate F&B imports used for final consumption.[11] The two series are highly correlated, and the scanner data pick up the large fall in imports over the crisis period.[12]

Why use scanner data rather than macroeconomic data for the purpose of this study? Scanner data allow us to measure expenditures on domestic and imported goods consistently within a single large dataset using final consumer prices. In contrast, macroeconomic data would require combining data on trade flows with household expenditure data, which would create multiple issues for the measurement of expenditures on domestic/foreign goods. One issue is that household expenditures are measured in final consumer prices and include domestic retail services, while trade flows are measured at the dock (Berger et al., 2009; Burstein et al., 2005). Another issue is that inventories can drive a wedge between final expenditures and trade flows, especially during sudden stop episodes (Alessandria et al., 2010). As a result, studies that use macro data have focused on relative price movements between domestic/imported goods, while with scanner data we can empirically examine both relative prices as well as *expenditure switching* between domestic and foreign goods and study the impact of relative price movements on consumption behavior.[13]

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[10]The CPI is constructed using multilateral GEKS price index. For an in-depth discussion of this index see e.g. Ivancic et al. (2011).

[11]We rely on the Global Trade Information Services (http://www.gtis.com), and the UN Broad Economic Classification (http://unstats.un.org/unsd/cr/registry/regcst.asp?Cl=10) in order to calculate the aggregate numbers.

[12]Aggregate F&B imports in customs data drop more quickly than in the store data and also show a more rapid recovery. This could be due to an inventory effect (e.g., such as argued by Alessandria et al., 2010). Though interesting for future research, this finding does not impact the analysis of the current paper given that we are interested in studying the total impact of the crisis, and not the dynamics per se.

[13]Another advantage is that scanner data report prices and quantities at the detailed item level, allowing

We should also note the limitations of the scanner dataset for studying expenditure switching. One obvious shortcoming is that the data only contain demand for domestic and imported F&B, though as noted above, our scanner data are representative of aggregate prices and expenditures on F&B, and consumption of F&B make up a significant portion of aggregate household expenditures in Latvia. We also unfortunately cannot match these data with supply-side data in order to capture the impact on exports. Thus, our results only speak to one facet of external sector adjustment – imports.

3 Empirical Findings

The Latvian economy experienced a sharp contraction during the sudden stop, and this contraction was felt across all sectors of the economy, including consumption in food and beverages. Figure 3 uses quarterly data to plot the year-on-year (y-o-y) log change in real aggregate food consumption in the scanner data.[14] The figure depicts a classic "boom-bust" episode, where consumption was growing before the crisis, at which point it experienced a substantial drop, bottoming out at -16% in real terms over Q4:08–Q4:09, which we define as our crisis period.

The scanner data allow us to document three empirical findings pertaining to expenditure switching during the crisis, with a focus on the relative movements of the domestic and foreign components of consumption and prices within and across narrowly defined product groups. These findings underpin the main results of the paper, as well as motivate the modeling and estimation methodology we use below.

The three empirical findings are:

1. Expenditure switching from imported to domestic food accounted for one third of the contraction of imports during the crisis, and was driven mainly by switching between items *within* narrowly defined product groups.

2. The expenditure switching was accompanied by a 4.4% rise in the relative price of foreign goods to total food CPI, where the relative price change was driven almost entirely by changes in prices *across* product groups.

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for a breakdown of relative prices and expenditures into narrow product groups of homogeneous items. Furthermore, scanner data directly record quantities purchased for each good, while National Income Account (NIA) data estimate quantities indirectly using surveys, which are bound to be less reliable.

[14]The sample begins at the second quarter of 2006, which is defined as May–July in order to maximize observations. Using y-o-y changes helps avoid seasonality issues.

3. Within the narrowly defined product groups, consumers systematically switched from higher unit value imported items to lower unit value domestic items during the crisis, which generated expenditure switching without any need for adjustment in relative prices.

3.1 Finding 1: Expenditure Switching

We first examine the role of expenditure switching in the total fall of imports during the crisis by considering a simple decomposition. We begin by defining X as total expenditures on F&B, and X^F as expenditures on imported F&B. Then, define $\Delta x \equiv \ln\left(X_{Q4:09}/X_{Q4:08}\right)$ and $\Delta x^F \equiv \ln\left(X_{Q4:09}^F/X_{Q4:08}^F\right)$ as the respective growth rates of expenditures. We can then decompose the fall in imports as

$$\Delta x^F \equiv \Delta x + \left(\Delta x^F - \Delta x\right),$$

where the first term on the right-hand side is the contribution due to an across-the-board contraction in food consumption, which is proportional to the fall in total consumption of F&B, and the second term is a residual that captures the contribution of expenditure switching. This term captures the finding that expenditures on imports contracted more than proportionally with aggregate expenditures on food, and as a result, there was expenditure switching from imported to domestic food. In the scanner data, imports fell by 26%, while total food expenditures fell by 18% during the crisis. Therefore, the expenditure switching term accounted for 8 percentage points, or one third, of the fall in imports.

Figure 4 provides an alternative way of quantifying the size of the expenditure switching, by plotting the y-o-y percentage point change in the import expenditure share, $\Delta(X^F/X)$. The solid line in the figure shows that at the trough (i.e., Q4:09), 3.5 percent of expenditures were reallocated from imports towards domestically produced food.

Although there is entry and exit of items in the scanner data, we find that the adjustment at the *intensive* margin accounts for the bulk of the expenditure switching in Latvia (See Appendix A for details). Given the relatively short horizon of our analysis, it is not that surprising that the extensive margin does not play a large role in the crisis dynamics, as, for example, inventories may have dampened the extensive margin supply response in the short-run. Furthermore, our findings are consistent with those of the recent trade collapse literature, which also finds that the extensive margin played a small role (see Bems et al., 2012, for a recent review), as well recent evidence on the import behavior of firms during the Argentinean crisis Gopinath and Neiman (Forthcoming, 2013).

We next exploit the data at both the product group and item level in order to distinguish between two sources of expenditure switching due to consumers reallocating expenditures either (i) *across* product groups, or (ii) between domestic and foreign items *within* product groups. The within margin can contribute directly to expenditure switching, as consumers substitute between similar domestic and foreign items. The across margin can contribute indirectly to expenditure switching as long as product groups have different import shares. For example, if the dairy product group is mainly composed of domestic items, while the alcohol product group has a large foreign content, then substitution from alcohol to dairy, holding all else equal, would result in aggregate expenditure switching.

Begin by defining a product group $g \in \{1, ..., G\}$, an item $i \in I_g$, and expenditure share s_{igt} for item i in product group g in period t, so that $\sum_g \sum_i s_{igt} = 1$. Further, denote $s_{gt}^j = \sum_{i \in I_{gt}^j} s_{igt}$ as the expenditure share for a subset j of items in product group g. With this notation one can express expenditure share of a product group as $s_{gt} = \sum_{i \in I_{gt}} s_{igt}$, and total expenditure share on imports is $s_t^F = \sum_g \sum_{i \in I_{gt}^F} s_{igt}$, where F refers to imported items.

Next, define the share of imports within a product group as $\varphi_{gt}^F = s_{gt}^F / s_{gt}$. Then $s_t^F = \sum_g s_{gt} \varphi_{gt}^F$, and aggregate expenditure switching between any two periods k and t can be decomposed into the two components of interest – expenditure switching within and across product groups – as follows:

$$
\begin{aligned}
s_t^F - s_k^F &= \sum_g s_{gt} \varphi_{gt}^F - \sum_g s_{gk} \varphi_{gk}^F \\
&= \underbrace{\sum_g s_{gk} \left(\varphi_{gt}^F - \varphi_{gk}^F \right)}_{\text{Within}} + \underbrace{\sum_g \varphi_{gk}^F \left(s_{gt} - s_{gk} \right)}_{\text{Across}} + \underbrace{\sum_g \left(\varphi_{gt}^F - \varphi_{gk}^F \right) \left(s_{gt} - s_{gk} \right)}_{\approx 0}. \quad (1)
\end{aligned}
$$

Figure 4 plots this decomposition for y-o-y changes in s_t^F, where a product group g is defined at the 4-digit level. We find that the bulk of expenditure switching took place within sectors (dash-dot line), as consumers substituted from foreign to domestic goods, while maintaining relatively constant shares of expenditures across product groups throughout the sample. The *within*-switching is a crucial empirical finding that our analysis incorporates below. To the best of our knowledge, there is no previous empirical work that either quantifies the magnitude of expenditure switching during a crisis, or decomposes expenditure switching within and across product groups. We therefore view these results as novel.

3.2 Finding 2: Relative Price Adjustment

We next examine price movements of domestic and imports goods at the aggregate and product group level. In order to do so, we must construct comparable price indexes across product groups from the UPC-level data on unit values and quantities. For our baseline results we construct aggregate prices using discrete Divisia (Törnqvist) price indexes.[15] The overall price index for F&B is

$$\Delta \ln P_t = \sum_g \sum_j \sum_{i \in I_{gt}^j} w_{igt} \Delta \ln p_{igt}, \tag{2}$$

where p_{igt} is the unit value of item i in product group g, $w_{igt} = 1/2 \left(s_{igt} + s_{igt-1} \right)$ is a corresponding expenditure-based weight, and $j = \{D, F\}$ sorts items by source (Domestic/Foreign) within each product group. Narrower price indexes of interest are computed as components of the overall price index. For example, price changes in product group g are

$$\Delta \ln P_{gt} = \frac{1}{\sum_j \sum_{i \in I_{gt}^j} w_{igt}} \sum_j \sum_{i \in I_{gt}^j} w_{igt} \Delta \ln p_{igt}, \tag{3}$$

and price changes for imported items in product group g are

$$\Delta \ln P_{gt}^F = \frac{1}{\sum_{i \in I_{gt}^F} w_{igt}} \sum_{i \in I_{gt}^F} w_{igt} \Delta \ln p_{igt}. \tag{4}$$

In order to link the relative price adjustment to our measure of expenditure switching, we define the aggregate relative price of imports as P_t^F / P_t, where P_t^F and P_t are, respectively, price indexes for aggregate imports and aggregate food consumption. The solid line in Figure 5 plots the y-o-y change in $\ln(P_t^F/P_t)$. The relative price increases by 4.4% y-o-y during the crisis period (Q4:08–Q4:09), and by 6% from trough to peak.

As with expenditure switching, it is instructive to decompose the change in the relative price into *across* and *within* product-group components. First, note that the (log) relative price can be written as a weighted sum of product-group relative prices:

$$\ln \frac{P_t^F}{P_t} = \sum_g \frac{w_{gt}^F}{w_t^F} \ln \frac{P_{gt}^F}{P_t} = \sum_g \frac{w_{gt}^F}{w_t^F} \left(\ln \frac{P_{gt}^F}{P_{gt}} + \ln \frac{P_{gt}}{P_t} \right),$$

where $w_{gt}^F / w_t^F = \sum_{i \in I_{gt}^F} w_{igt} / (\sum_g \sum_{i \in I_{gt}^F} w_{igt})$ is the import share of group g in total imports. We can then express the growth rate of the relative price between periods k and t

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[15]Findings of this paper are robust to the use of alternative price index definitions, such as a multilateral GEKS price index or a Fisher price index, in construction of aggregate prices.

as

$$
\begin{aligned}
\ln \frac{P_t^F}{P_t} - \ln \frac{P_k^F}{P_k} &= \sum_g \frac{w_{gt}^F}{w_t^F}\left(\ln \frac{P_{gt}^F}{P_{gt}} + \ln \frac{P_{gt}}{P_t}\right) - \sum_g \frac{w_{gk}^F}{w_k^F}\left(\ln \frac{P_{gk}^F}{P_{gk}} + \ln \frac{P_{gk}}{P_k}\right) \\
&= \underbrace{\sum_g \frac{w_{gk}^F}{w_k^F}\left(\ln \frac{P_{gt}^F}{P_{gt}} - \ln \frac{P_{gk}^F}{P_{gk}}\right)}_{\text{Within}} + \underbrace{\sum_g \frac{w_{gk}^F}{w_k^F}\left(\ln \frac{P_{gt}}{P_t} - \ln \frac{P_{gk}}{P_k}\right)}_{\text{Across}} \\
&\quad + \underbrace{\sum_g \left(\frac{w_{gt}^F}{w_t^F} - \frac{w_{gk}^F}{w_k^F}\right) \ln \frac{P_k^F}{P_k}}_{\approx 0} .
\end{aligned}
\tag{5}
$$

In Figure 5, again using 4-digit product groups, one can see that the increase in the relative price of imports was almost exclusively driven by price movements *across* product groups (dash line).[16] Within product groups (dash-dot line), relative prices did not exhibit any systematic deviations. This result is the opposite of what occurred for expenditure shares, where switching took place within, not across product groups. From the conventional macroeconomic theory standpoint these findings present a puzzle: why are consumers switching expenditures towards domestic items within product groups, if domestic items are not becoming relatively less expensive than their foreign counterparts?

Our empirical findings about the price adjustment can be related to extensive macroeconomic literatures on relative price adjustment during crises/sudden stops (e.g., see Burstein et al., 2005; Kehoe and Ruhl, 2008; Mendoza, 2005; Obstfeld and Rogoff, 2005). Appendix B presents a methodology for mapping the micro scanner data into aggregate tradable/nontradable goods based on items' *shelf-life*, and provides a decomposition of the adjustment in the relative price of imports into contributions from *internal* – nontradables/tradables – and *external* – domestic/imported tradables – margins. We can then reinterpret Finding 2 in a conventional two-sector macroeconomic framework. The magnitude of the overall import price adjustment in Latvia (i.e., 6%) was a fraction of comparable price adjustments in earlier episodes (e.g., see Burstein et al., 2005). This is not surprising, given that the earlier episodes were accompanied by large nominal devaluations, while Latvia maintained a peg to the Euro. Ignoring the difference in the overall magnitude of the relative price change compared to other crisis episodes, we find that the bulk of the relative price adjustment during the crisis took place between the relative price of tradables

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[16]Given some policy actions taken by the government during the crisis, such as increase in taxes on alcoholic beverages, we have examined whether any given product group drove this change in relative price. We found that no single product group drove the movement in the aggregate relative price, including alcoholic beverages.

and nontradables (i.e., a change along the internal margin) which is consistent with findings in earlier empirical work on sudden stop adjustments.

3.3 Finding 3: Shifts in Within-Group Item Mix

In the absence of relative price adjustment within sectors, this section examines other possible sources of expenditure switching within product groups.

3.3.1 Differences in Item Unit Values

We start by investigating whether there are systematic differences in unit values across comparable UPC items, as well as between comparable domestic and imported items, by examining unit value differences within detailed 4-digit product groups.

In order to document whether such differences exist, we restrict the data to comparable 'net' units (e.g., kilograms, liters) within product groups. We first drop product groups where 'pieces' are used as the measure of units, because such units are not comparable across items. This leads to the dataset's total revenues dropping by 7.6%. Next, we identify the most frequent units for each product group and drop items that are not measured in such units (for example, some product groups might report both L and KG unit values). This decreases total revenues by a further 2.1%.

There is a great deal of heterogeneity in the dispersion of unit values across product groups. Figure 6 plots the distribution of interquartile ranges of unit values for each product group, where the interquartile range of a given product group is defined as the difference between the unit value of the goods at the 75th and 25th percentiles of the product group's distribution of unit values. We find that for the median product group, the unit value at the 75th percentile is 73% above that of the 25th percentile.

The observed scope of unit value dispersion within product groups suggests that expenditure switching could be driven by substitution from high value to low value items, irrespective of relative price changes. To have a better view on this, we next calculate unit values for domestic and imported components for the quarters for which data are available.[17] In particular, we compute the domestic and import unit values of product group g, V_{gt}^j as

$$V_{gt}^j = \frac{\sum_{i \in I_{gt}^j} p_{igt} q_{igt}}{\sum_{i \in I_{gt}^j} q_{igt}} = \sum_{i \in I_{gt}^j} \phi_{igt} p_{igt}, \tag{6}$$

---

[17]For the domestic-foreign unit value comparison we identify product groups that account for at least 0.01% of total revenues and where both domestic and imported components account for at least 5% of the product group's expenditures. When doing so, we are left with a sample of 265 4-digit product groups.

where $j = \{D, F\}$, $\phi_{igt} = q_{igt} / \sum_{i \in I_{gt}^j} q_{igt}$ is a quantity-based weight, and p_{igt} is the unit value for item $i \in I_{gt}^j$.

Figure 7 plots the distribution of the resulting unit value differences. The within product groups unit value of the imported component is on average 33% higher than that of the domestic component. The median difference is 30%. These differences are persistent over time, varying in the 27–37% range. Therefore, over the whole sample, foreign goods are on average more expensive than comparable domestic ones.

3.3.2 Flight to Cheaper Substitutes

To further investigate unit value differences within product groups as a source of expenditure switching, we next turn to a comparison of within product group price indexes and average unit values. Following Boorstein and Feenstra (1987), we compute a change in product group's g average unit value between t and $t-1$ as

$$\Delta \ln V_{gt} = \ln \sum\nolimits_{i \in I_{gt/t-1}} \phi_{igt} p_{igt} - \ln \sum\nolimits_{i \in I_{gt/t-1}} \phi_{igt-1} p_{igt-1}, \tag{7}$$

where attention is restricted to continuing items.[18] The corresponding change in the price index is

$$\Delta \ln P_{gt} = \frac{1}{w_{gt}} \sum\nolimits_{i \in I_{gt/t-1}} w_{igt} \left(\ln p_{igt} - \ln p_{igt-1} \right), \tag{8}$$

and an index that captures deviations between (7) and (8) is then defined as

$$\Delta \ln W_{gt} = \Delta \ln V_{gt} - \Delta \ln P_{gt}. \tag{9}$$

The key source of deviations between the average unit value and the price index is that the price index aggregates changes in item unit values for some fixed vector of item weights. In this case, a shift in the consumed item mix, $\Delta q_{igt} \gtrless 0$, on its own cannot generate any changes in the aggregate price index. In contrast, a group's average unit value takes into account changes in *both* item prices and quantities consumed. $\Delta \ln W_{gt}$ therefore captures changes in the mix of consumed items within product groups.

Of particular interest to us is the case when the item mix in a product group shifts systematically towards lower unit value items. In this case, a group's average unit value falls, $\Delta \ln V_{gt} < 0$, but the price index remains constant, $\Delta \ln P_{gt} = 0$, resulting in $\Delta \ln W_{gt} < 0$. The opposite holds if the consumed item mix shifts towards more expensive items.

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[18]The extensive margin (i.e., entering and exiting items) contributes an upward trend to groups' average unit value over the sample. This trend reflects the common practice of stores hiding price increases via the introduction of new items to the shelf. However, importantly for our purpose, the contribution of the extensive margin did not vary systematically with the crisis.

We apply this measure of the shift in item mix to our dataset. For each 4-digit product group g and each t we compute ΔW_{gt} based on quarter-on-quarter (q-o-q) data. We then aggregate the resulting index across product groups using expenditure weights. Finally, the aggregate index is cumulated over the whole sample period and expressed in y-o-y terms to eliminate seasonality.

The solid line in Figure 8 shows that the item mix within product groups varied systematically with the boom-bust cycle. In particular, during the crisis the average unit value fell relative to the price index, indicating that consumption shifted systematically towards items with lower unit values. This finding holds in aggregate data as well as within items of imported and domestic origin.

The key question for this paper is whether the shift in item mix towards lower unit values during the crisis, as documented in Figure 8, also induced expenditure switching as consumers switched from expensive foreign items to cheaper domestic ones. To capture such a link, we substitute $\Delta \ln P_{gt}$ in (8) with a partial price index, which imposes item homogeneity across domestic and imported items in each product group. A product group's aggregate price index can then be computed by first deriving the average unit value within a group for continuing domestic and imported items:

$$\Delta \ln V_{gt}^{j} = \ln \sum\nolimits_{i \in I_{gt/t-1}^{j}} \left(\phi_{igt} p_{igt} - \phi_{igt-1} p_{igt-1} \right), \tag{10}$$

and then computing a partial price index, $P_{gt}^{F/D}$, based on the two unit values series, V_{gt}^{F} and V_{gt}^{D}:

$$\Delta \ln P_{gt}^{F/D} = \sum\nolimits_{j=F,D} w_{gt}^{j} \Delta \ln V_{gt}^{j}. \tag{11}$$

This partial price index eliminates all possible sources for deviations between the average unit value and full price index, except changes in the product mix that stem from a shift in consumption between domestic and imported items. The resulting index,

$$\Delta \ln W_{gt}^{F/D} = \Delta \ln V_{gt} - \Delta \ln P_{gt}^{F/D}, \tag{12}$$

is by construction zero for product groups that contain only domestic or only foreign items. It is also zero in product groups that contain items from both sources, but where average unit values do not differ across sources, i.e., $V_{gt}^{F} = V_{gt}^{D}$. Only if both (i) imported items exhibit higher average unit values and (ii) consumption during the crisis shifts towards items with lower unit values, does this modified index take negative values.

The dashed line in Figure 8 plots this modified index. We find that during the crisis there was a systematic shift in the consumption mix towards domestic items, which exhibited

lower unit values. As expected, the size of the modified index, $\Delta \ln W_{gt}^{F/D}$, is a fraction of the full index, $\Delta \ln W_{gt}$, as heterogeneity at the item level is reduced to two aggregated items (domestic and imported). Nevertheless, the partial index shows that during the crisis consumers systematically switched from higher unit value imported items to lower unit value domestic items.

Our finding that there is switching towards items with lower unit values during crises is consistent with earlier findings by Burstein et al. (2005). To the best of our knowledge, however, there is no evidence on differences in the unit values of domestic and imported goods given the lack of available data, nor on how consumers switch from higher unit value foreign goods to lower unit value domestic goods. Work examining scanner data in the U.S. has noted that consumers search for cheaper goods by switching stores during recessions (Coibion et al., 2012), as well as differences in consumption across cities and household income levels (Handbury, 2012), but no one has examined the international dimension yet. The finding that imported goods tend to be more expensive than domestic ones in Latvia is an important result, which we build on below in helping to explain the observed expenditure switching.

4 Model

To formally quantify the importance of relative prices and income on expenditure switching, we next model the consumer's expenditure allocation for F&B. The conventional approach in the international macroeconomics literature is to explicitly model the consumer's choice between goods of domestic and foreign origin in a model that also distinguishes between tradable and nontradable goods (see Obstfeld and Rogoff, 2005, for example). However, given the item-level data at our disposal, we find it more intuitive to model the consumer as basing her consumption decisions on the characteristics of goods, such quality or potential calories per unit (since we are considering food). These characteristics in turn may be reflected in a product's relative price or unit value. *A priori*, it is not clear why a consumer would explicitly discriminate between geographical origin of a good given these other characteristics.

We therefore follow the literature that uses scanner data and model the expenditure allocation as a two-stage decision, where a consumer first allocates expenditures *across* grocery product groups (tea, coffee, cacao, etc), and then allocates expenditures between UPC items *within* product groups.[19] Given the documented heterogeneity in unit values

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[19]See, Broda and Weinstein (2010) or Handbury (2012) for recent contributions using nested utilities and

within product groups (see Finding 3), we also build in a channel through which consumers may substitute between low and high priced goods when faced with an income shock, such as the one experienced by Latvia during the crisis. In particular, we borrow from the setup of Hallak (2006)'s model, which allows goods to vary by quality, and for the consumer's intensity of demand for quality to depend on her income level. These modifications of the standard CES demand system introduce a non-homotheticity at the bottom layer of the utility function.[20] The higher the income the more the consumer values the higher quality items. Though we are not modeling quality formally here, since other factors may drive the difference in prices across domestic and foreign goods (e.g., transport costs), this modeling strategy captures an important potential channel that we wish to test; i.e., that consumers substituted to cheaper goods during the crisis, irrespective of relative price changes.

Introducing a "quality" parameter into the model that allows for non-homothetic preferences to play a role in expenditure switching is novel, and has not been explored in the international macroeconomics literature in general. The few applications of non-homothetic preferences in macroeconomics usually rely on Stone-Geary type utility functions (Herrendorf et al., Forthcoming, 2013; Kongsamut et al., 2001; Ravn et al., 2008).

4.1 Setup

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

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

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

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

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scanner data. Blackorby et al. (1978) is an early contribution that uses nested utility, and which also allows for non-homothetic preferences.

[20]Hallak (2006) takes the supply of quality and income as exogenous in a partial equilibrium setting, like ours. See Feenstra and Romalis (2013) for a general equilibrium model, where quality is an endogenous outcome. Furthermore, Choi et al. (2009) and Fajgelbaum et al. (2011) study how countries' income distributions affects trade and quality in a more general setting.

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.

4.2 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's hold both at the top and bottom levels of the utility. Specifically, at the top level we have

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

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

$$s_{gt} \equiv \frac{P_{gt}c_{gt}}{C_t} = \omega_{gt}P_{gt}^{1-\rho}. \tag{13}$$

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

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

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

$$\varphi_{igt} \equiv \frac{p_{igt}c_{igt}}{P_{gt}c_{gt}} = \frac{1}{N_{gt}} \left(\frac{\frac{p_{igt}}{\theta_{ig}^{\lambda_g(C_t)}}}{P_{gt}} \right)^{1-\sigma_g}, \tag{14}$$

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

$$s_{igt} \equiv \varphi_{igt}s_{gt} = \frac{1}{N_{gt}} \left(\frac{\frac{p_{igt}}{\theta_{ig}^{\lambda_g(C_t)}}}{P_{gt}} \right)^{1-\sigma_g} \omega_{gt}P_{gt}^{1-\rho}. \tag{15}$$

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

$$P_{gt} = \left(\frac{1}{N_{gt}} \sum_i \left(\frac{p_{igt}}{\theta_{ig}^{\lambda_g(C_t)}} \right)^{1-\sigma_g} \right)^{\frac{1}{1-\sigma_g}}. \tag{16}$$

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

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

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}.*

5   Estimation and Results

This section uses the model presented above to develop an estimation strategy, which exploits the richness of the item-level data in a panel regression framework. We first show how to arrive at the estimating equation using the theoretical expenditure model, where we focus on exploiting the within-product group variation of the data. Then, given the estimated parameters, we predict an item's share in a given product group, and aggregate all predicted shares of imported goods in order to calculate the predicted import share of total F&B, and the corresponding expenditure switching over periods, which we compare to the data.

5.1 Setting up the Estimation Equation

The key equation that characterizes the solution of the model presented in the previous section is (15). 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.[21] We will discuss these issues in detail below.

First, log-differencing (15) and substituting in (13), we arrive at

$$\Delta \ln \varphi_{igt} = \Delta \ln N_{gt} + (1 - \sigma_g)\Delta \ln \left(\frac{p_{igt}}{P_{gt}} \right) + (\sigma_g - 1)\Delta\lambda_g(C_t) \ln \theta_{ig}. \tag{17}$$

To allow for estimation of (17), 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 (17) as

$$\Delta \ln \varphi_{igt} = \Delta \ln N_{gt} + (1 - \sigma_g)\Delta \ln \left(\frac{p_{igt}}{P_{gt}} \right) + (\sigma_g - 1)\mu_g \ln \theta_{ig}\Delta \ln C_t, \tag{18}$$

where the η_g disappears from taking first differences,[22] and since the aggregate price index, P_t, is the numéraire, C_t is expressed in real terms.

5.2 Taking the Model to the Data

The estimating equation (18) is based on a partial equilibrium model, which treats several variables (such as the number of items in a group at a given time, N_{gt}) as given, and ignores potential supply-side shocks that may impact both the quantity and the price of goods. Furthermore, the aggregate price indexes are model-based, and therefore a function of some of the parameters we wish to estimate. We must address these issues in order to insure we have sufficient data to estimate the model, and to obtain consistent estimates.

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

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[21]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.

[22]Eliminating $\eta_g \times \ln \theta_{ig}$ by first differencing the log shares is helpful, since we would otherwise have to include item-level fixed effects in the panel regressions below, which would make estimation infeasible given the extremely large number of items in the dataset, along with a relatively short time series.

the Törnqvist index. This approach may lead to measurement error due to unaccounted for income-driven substitution, which would be picked up by the model-based group price index of (16). However, since $\Delta \ln P_{gt}$ enters the estimating equation linearly (both in the relative price and in deflating total expenditures), we can 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 et al., 2011). Moreover, some firms (either domestic or foreign) may have been driven out of business, thereby impacting the price level and supply of goods in a given product group. Again, the inclusion of the product group×time fixed effects will control for these potential shocks at a very disaggregated level, and capture the general equilibrium impact of the shocks *within* a product group.

Finally, arguably reverse causality could bias our results. For example, faced with a large demand shock arising from the crisis, Latvian consumers may have become more nationalistic and thus switched to consuming more domestic items. This switch would put upward pressure on domestic prices viz. foreign ones, thus generating a *negative* relationship between relative prices and import shares. However, we do not observe such price or consumption behavior in the data. Moreover, including group×time fixed effects will help control for this potential nationalism biasing the estimation of how changes in income affects changes in expenditure shares.

We therefore proceed by writing the estimating equation for (18) as

$$\Delta \ln \varphi_{igt} = \alpha_{gt} + \beta_{1g} \ln \Delta \left(\frac{p_{igt}}{P_{gt}} \right) + \beta_{2g} \ln \bar{p}_{ig} \Delta \ln C_t + \varepsilon_{igt}, \tag{19}$$

where α_{gt} is a 4-digit product group×time fixed effect, which absorbs all explanatory variables that only vary in the gt dimension; $\beta_{1g} = 1 - \sigma_g$; $\beta_{2g} = \mu_g(\sigma_g - 1)$; $\bar{p}_{ig}$ is a proxy for the quality parameter θ_{ig}, and is calculated as the sample median of each item's relative unit value standing within a product group, p_{igt}/V_{gt}, where a group's average unit value V_{gt} is defined in (6), and ε_{igt} is as random disturbance term. As in the model, we interact the "quality" term $\ln \bar{p}_{it}$ with the growth rate of income, $\Delta \ln C_t$. We use quarterly real per-capita household consumption for C_t, which we take from the International Financial Statistics (IMF).

The inclusion of product group×time fixed effects implies that the β parameters will be identified from variation across items within their product groups during a given period. This within variation is crucial for the identification of β_{2g}, since the unit value of a given

item ($\bar{p}_{ig}$) is only comparable within a group.

In order to estimate (19), we use the same data sample as used in Section 3, though we drop four 4-digit product groups given that they do not contain enough data to identify the coefficients of interest, and we do not trim the data based on import shares as we did in Section 3.3. The final regression sample comprises of 236,519 item×time observation, and 384 product groups.[23]

We estimate two versions of the baseline regression (19). The first model, which we call the 'CES' model, restricts all β_{2g} to 0. Therefore, only changes in relative prices will affect an item's share. The second model, which we call the 'NH' (non-homothetic) model, runs (19) unrestricted. Table 3 presents summary statistics of the distribution of the estimated βs for each product group for the two models. Column (1) presents the CES model, where we restrict β_{2g} to be 0. The median value of β_{1g} is -1.925, implying a price elasticity, σ_g, equal to 2.925. There is quite a bit of dispersion in the estimated coefficients, as well as their standard errors,[24] but in total, 273 coefficients are significant at the 10% level or below. Columns (2) and (3) present summary statistics for the β_{1g}s and β_{2g}s, respectively, for the unrestricted model. The median value of β_{1g} is slightly larger (in absolute value) than that of the restricted model, with a value -1.955. The estimated income coefficients, β_{2g}, are on average positive, as we would expect, with a median value of 0.968. The median quality measure across all groups is 0.1583, which implies a median value of μ_g equal to 3.127. Again, there is considerable heterogeneity in the estimated coefficients across groups, but the majority of the price coefficients are significant at the 10% level or below, as are one-third of the income coefficients. The R^2 is somewhat larger for the NH model (0.103 vs. 0.099), and a chi-square test rejects no systematic difference in the coefficients of the two models.

Table 4 presents summary statistics for only the estimated coefficients that are significant in our baseline regressions. Column (1) presents the summary statistics for 270 of the relative price coefficients that are reported in Column (1) of Table 3, and that are significant. The median value of the coefficient is larger, in absolute value, for the sub-sample of significant coefficients (-2.435 vs. -1.925). We find a similar results when comparing median value of the significant relevant price coefficients for the NH model in Column (2). The 273 coefficients have a median value of -2.416, which is larger than median value across

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[23]We also experimented with restricting the sample so that each product groups contains a minimum of 500 observations over the whole sample period, which cuts the sample to 143 product groups. Furthemore, we also run regressions dropping the alcoholic beverage product groups. Results were robust to these restrictions, and are available upon request.

[24]Not reported to conserve space, but available from the authors upon request.

all coefficients (-1.955) reported in our baseline results. Finally, the median of the 101 significant coefficients reported in Column (3) is over three times larger than the median value of all the coefficients, which is reported in Table 3: 3.347 vs. 0.968 (of the 101 coefficients, 80 are indeed positive).[25]

We also consider some additional robustness checks. First, we allow for the possibility of a non-linear effect by including a squared term of the change in aggregate expenditures, interacted with quality. Columns (1)-(3) of Table 5 present the summary statistics of the estimated coefficients for the relative price, income effect, and income effect squared, respectively. The median coefficients for the relative price and income effect barely change relative to the baseline results in Table 3, while the median coefficient value for the squared income term is zero. Therefore, it does not appear that non-linearities are a concern.

Second, we allow for the possibility that domestic suppliers reacted differently than foreign ones during the boom and following crisis. For example, data show that producers in Latvia responded to the severe crisis by cutting production costs (e.g., wages), which could lower prices of domestic final goods, including food items, relative to their imported counterparts (Blanchard et al., 2013; Kang and Shambaugh, 2013). To investigate this possibility, we interact the product group×time effect with a domestic/foreign dummy variable. Columns (4)-(6) of Table 5 present the summary statistics of the estimated coefficients for the CES and NH specifications. The median values for the coefficients for the relative prices do not vary dramatically relative the baseline results of Table 5, while the median coefficient on the income effect drops slightly (by 0.1).[26] Finally, we also investigate the quantitative impact on our estimated coefficients of not including product group×time fixed effects. Table A2 presents estimates for the CES and NH models for an array of fixed effects configurations, including none. Moving from left to right (Columns 1-4), we note that, for both models, the median value of the relative price coefficient increases in absolute value when including more fixed effects. Magnitudes of the coefficients are smaller (in absolute value) than those in the preferred specifications in Table 3. Therefore, controlling for product-group and time effects appears to help deal with potential attenuation bias from mis-measurement of price indexes.

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[25]We also run a "top layer" regression at the product group level to estimate ρ. We regress the log change of product group expenditure shares, $\Delta \ln s_{gt}$, on the log change of relative prices, $\Delta \ln(P_{gt}/P_t)$. Table A1 reports the results for different arrays of fixed effects. All point estimates are significant at the 1% level, and the implied value of ρ, for the most stringent set of fixed effects (group and time), is 2.26.

[26]Note that a drop in prices of the domestic non-tradable retail services is not a concern for us, as long as the fall in retail margins is applied proportionally to domestic and imported items within narrow product groups. Furthermore, the fixed effects also capture any commonality of this phenomena across items.

5.3 Predicted Within-Product Group Expenditure Switching

Before moving on to a full-scale exercise in order to predict aggregate expenditure switching, we would like to have an idea of the magnitude of the expenditure switching observed within an average product group in the economy over the crisis period, and how this matches up with what is predicted from using the estimated price and income coefficients from the baseline regressions in Table 3.

Since we will calculate the predicted growth rate of shares excluding the group×time fixed effect, we first demean the actual growth rates observed in the data by their group-time mean for comparability. Over the four quarters of the crisis, the resulting average growth rate of the import share across product groups observed in the data is -0.067. Over the same period, the average growth rate of relative prices $(\Delta \ln(p_{igt}/P_{gt}))$ for foreign items across product groups is -0.0003, and the growth rate of real per-capita consumption is -0.080. Next, the measure we use as a proxy for quality in the regression $(\ln \bar{p}_{ig})$ has an average value of 0.304 for foreign items entering the regression during the crisis period (the average over the whole sample is 0.29). Armed with average values of $\Delta \ln(p_{igt}/P_{gt})$ and $\ln \bar{p}_{ig}$ for imported items as well as the average value of $\Delta \ln C_t$, we can then calculate a predicted growth rate of a foreign item's share within a product group. To do so, we use the median estimated coefficients in Columns (2) and (3) in Table 3.[27] Applying these coefficients to the data moments yields a predicted growth rate of the import share of -0.018, or 27% of what is observed in the data. The relative price component of the prediction is almost zero (0.005), and the income component of the prediction is -0.023. Therefore, and as we shall see below, the income effect plays an important role in driving expenditure switching within product groups.

Given the disaggregated nature of the data, and that we are running regressions using growth rates, explaining 27% of the observed expenditure switching within an "average" product group over the crisis is non-trivial. However, there is substantial variation in the estimated coefficients across product groups, as well as heterogeneity in how import intensive each group is. Therefore, to see how well the regression estimates perform in predicting aggregate within expenditure switching, we next allow for the heterogeneity in coefficients, relative price changes, and import shares observed in the data.

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[27]The mean coefficient value for the price coefficient is -1.91, and the mean income coefficient is 2.80. Therefore, we may be understating the impact of the income effect in this simple exercise.

5.4 Predicted Aggregate Expenditure Switching

We next predict the *aggregate* within-group y-on-y expenditure switching. We start by calculating predicted expenditure switching between any consecutive quarters τ and $\tau - 1$:

$$(\widehat{s_\tau^F - s_{\tau-1}^F})^{\text{Within}} = \sum_g s_{g\tau-1}(\widehat{\varphi}_{g\tau}^F - \varphi_{g\tau-1}^F), \tag{20}$$

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

1. Take the estimated coefficients from the within-group regressions (19), $\widehat{\beta}_{1g}$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 $\Rightarrow \widehat{\varphi}_{g\tau}^F = \sum_{i \in I_{g,\tau}^F} \widehat{\varphi}_{ig\tau}$.

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:

$$(\widehat{s_t^F - s_k^F})^{\text{Within}} = \sum_{\tau=k}^{t} \sum_g s_{g\tau-1}(\widehat{\varphi}_{g\tau}^F - \varphi_{g\tau-1}^F), \tag{21}$$

where $k = t - 3$. Appendix C provides more details on these steps, 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.[28]

We focus on (21) because, as Finding 1 above shows, the majority of expenditure switching occurred within product groups. Furthermore, we only look at the share changes as predicted by either (i) changes in relative prices, or (ii) the income effect, and thus do not include the group×time fixed effects in calculating the predicted growth rates of the items' shares.

Figure 9 plots the actual and predicted within-group y-o-y expenditure switching. The first fact to note is that the within-group expenditure switching observed in the data (the solid line) has very similar dynamics compared to the within-group component plotted in

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[28]This measure is not identical to the within expenditure switching presented in Figure 4, but the two series are very similar.

Figure 4, which was calculated using a slightly larger data sample (including items that did not exist for two consecutive periods). Next, comparing the two models' predicted expenditure switching to that of the data in Figure 9, the predicted expenditure switching for the CES model (dash-dot line) does a very poor job in tracking the actual within-group expenditure switching observed in the data, particularly during the crisis period when income dropped substantially. However, one sees that the non-homothetic model's predicted values (dash line) appear to track the data better throughout the sample, and matches the switching during the pre-crisis boom period, as well as the switching during the crisis. In particular, the within component of expenditure switching between Q4:08–Q4:09 implied by the data is -0.02, while the CES model predicts a value of only 0.001, and the non-homothetic model predicts a value of -0.016. Therefore, the CES model does not explain any expenditure switching between domestic and imported goods at the aggregate level (and in fact goes in the wrong direction), while the non-homothetic model is able to explain a little over 80% of what is observed in the data during the crisis period.

Finally, Figure 10 decomposes predicted expenditure switching of the non-homothetic model into separate price (dash-dot line) and income (dash line) effects. We calculate these by either shutting down the price effect ($\beta_{1g}s = 0$), or the income effect ($\beta_{2g}s = 0$), and then predict the log change in item's shares, and aggregate up. It is clear that the income effect is responsible for almost all of the predicted expenditure switching.

6   Conclusion

This paper measures what drove expenditure switching in Latvia during a sudden stop episode in 2008–09, using a scanner-level dataset of food and beverages. Contrary to conventional theory, relative price changes did not drive expenditure switching. Instead, this paper's findings show that the fall in income during the crisis led consumers to substitute from foreign to domestic goods, since foreign goods were on average more expensive than domestic ones. This non-homothetic channel is estimated using a simple model that allows for quality differences across goods, where the consumer's intensity of demand for quality varies with income.

The analysis in this paper only focuses on substitution between domestic goods and imports for a particular sector of the economy for a country that maintained a peg during its crisis. Future work should investigate how relevant the non-homothetic channel is in a more general setting, which incorporates exports and other sectors of the economy, as well studying as how results vary across exchange rate regimes.

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 et al. (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 Sections 4 and 5 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 et al. (2012), 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:

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

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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.[29] 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:

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

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

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

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

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

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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 3.5% 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   Relative Price Adjustment with Tradable and Nontradable Goods

The modeling and quantitative approach of the paper focuses on the detailed micro-level data in order to identify the margins at which expenditure switching took place, and then aggregate the results to draw conclusions about expenditure switching at the macroeconomic level. One downside of this micro-approach in studying the crisis/sudden stop is that our findings cannot be easily compared to the extensive existing literature (e.g., see Burstein et al., 2005; Kehoe and Ruhl, 2008; Mendoza, 2005; Obstfeld and Rogoff, 2005) that emphasizes the distinct roles of tradable and nontradable goods during the crisis.

To facilitate such a comparison, this appendix recasts Finding 2 of Section 3 from the perspective of a two-sector economy with tradable and nontradable goods. In order to do so, we must define whether a given 4-digit product group belongs to the tradable or nontradable sector. Though F&B as a whole is commonly considered a tradable sector in the international macroeconomic literature (Berka and Devereux, 2011; Crucini et al., 2005), Table 2 reveals that many product groups within F&B have a negligible import content.

We aggregate F&B product groups into tradables/nontradables based on the *shelf-life* of a typical product in each 4-digit product group (Boyer and McKinney, 2013). A 4-digit product group is defined as nontradable, if its shelf-life is less than 180 days and as tradable if the shelf-life is equal to or exceeds 180 days. With this classification 40% of expenditures on F&B are tagged as nontradables and remaining 60% of expenditures are classified as tradables, with imports constituting 55% of tradables. Of all imported food 90% fall into the tradable category.[30]

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[30]An alternative commonly pursued approach in the literature to distinguish between tradables/nontradables is to label a product group as tradables, if expenditures on imports in the group exceed 10%. Decomposition results for this approach were similar to the 'shelf-life' approach. However, we find import intensity to be a less appealing measure, as low import content might merely signal a comparative

We re-apply the price decomposition formula from Section 3.2 to the two-sector case with tradables and nontradables, i.e., $g \in \{N, T\}$. Adjustments within and across sectors can now be cast in more familiar terms:

$$\ln \frac{P_t^F}{P_t} - \ln \frac{P_k^F}{P_k} = \underbrace{\frac{w_{Tk}^F}{w_k^F} \left(\ln \frac{P_{Tt}^F}{P_{Tt}} - \ln \frac{P_{Tk}^F}{P_{Tk}} \right)}_{\text{External margin}} + \underbrace{\frac{w_{Tk}^F}{w_k^F} \left(\ln \frac{P_{Tt}}{P_t} - \ln \frac{P_{Tk}}{P_k} \right)}_{\text{Internal Margin}}$$

$$+ \underbrace{\frac{w_{Nk}^F}{w_k^F} \left(\ln \frac{P_{Nt}^F}{P_{Nt}} - \ln \frac{P_{Nk}^F}{P_{Nk}} \right)}_{\approx 0} + \underbrace{\frac{w_{Nk}^F}{w_k^F} \left(\ln \frac{P_{Nt}}{P_t} - \ln \frac{P_{Nk}}{P_k} \right)}_{\approx 0} \quad \text{(B.1)}$$

$$+ \underbrace{\sum_g \left(\frac{w_{gt}^F}{w_t^F} - \frac{w_{gk}^F}{w_k^F} \right) \ln \frac{P_k^F}{P_k}}_{\approx 0} .$$

Price adjustment *within* sectors amounts to the adjustment of the relative price of imports within the tradable sector, commonly referred to in the macro literature as the *external margin*. This is the case because the import content in the nontradable sector is negligible. Price adjustment *across* sectors captures changes in the relative price of tradables, or the *internal margin*.

Figure A3 plots the contribution of external and internal margins of the adjustment in the relative price of imports. The last three terms on the right hand side of (B.1) are collected in the residual term, which, as expected, is close to zero. The price decomposition results are broadly the same as at the level of the 4-digit product groups, presented in Figure 5, and the relative price adjustment during the crisis took place almost entirely at the internal margin, i.e., between the relative price of nontradables and tradables.[31] This result is consistent with empirical findings from other crisis episodes (e.g., see Burstein et al., 2005; Kehoe and Ruhl, 2008; Mendoza, 2005; Obstfeld and Rogoff, 2005), where authors find that the relative price of nontradables played a key role in the price adjustment.

Zooming in on the sources of the price adjustment between nontradables and tradables, we find that during the crisis price of tradables increased by 2.8%, while prices of nontradables decreased by 10.3%. Thus, the bulk of the adjustment took place through *deflation* in the nontradable sector. This finding contrasts earlier crisis episodes, which, accompanied by

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advantage of the domestic producers.

[31]Our results are robust to other threshold values for defining nontradables. Under an alternative specification a product group is defined as nontradables if its shelf-life is less than 60 days. In this case, nontradables constitute 28% of F&B expenditures and 96% of imports fall into the tradable category. The internal margin remains the main contributor to the adjustment in the relative price of imports, accounting for 3/4 of the overall price adjustment.

large nominal devaluations, exhibited rapid inflation in the tradable sector and considerably more sluggish price increases for nontradables (see Burstein et al., 2005).

An important implication of our Engel (1999)-type decomposition is that during the sudden stop episode in Latvia the relative price of imported tradables to all tradables, P_T^F/P_T, increased by a mere 0.3% (over Q4:09/Q4:08) in the F&B sector. For this price adjustment to explain the observed expenditure switching from foreign to domestic tradables, the price elasticity would have to be 25.6, which is a very large value for this level of aggregation. Therefore, even at this aggregate level of analysis, a further channel is needed to explain the observed expenditure switching. The introduction of a non-homothetic channel, like in the model of Section 4, into macroeconomic models could be one potential way to help explain the observed expenditure switching in a standard two-sector macroeconomic framework.

Appendix C   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 (19). 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.

C.1   Step 1

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

$$\widehat{\Delta \ln \varphi}_{ig\tau} = \widehat{\beta}_g Z_{ig\tau}, \tag{C.1}$$

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

C.2   Step 2

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

$$\widehat{\varphi}_{ig\tau} = \left(1 + \widehat{\beta}_g Z_{ig\tau}\right) \varphi_{ig\tau-1}. \tag{C.2}$$

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

C.3 Step 3

We next simply aggregate (C.2) for each group g for only foreign items to obtain a product group's predicted foreign share:

$$
\begin{aligned}
\widehat{\varphi}_{g\tau}^F &= \sum_{i \in I_{g\tau/\tau-1}^F} \left(1 + \widehat{\beta}_g Z_{ig\tau}\right) \varphi_{ig\tau-1} \\
&= \sum_{i \in I_{g\tau/\tau-1}^F} \varphi_{ig\tau-1} + \sum_{i \in I_{g\tau/\tau-1}^F} \left(\widehat{\beta}_g Z_{ig\tau}\right) \varphi_{ig\tau-1} \\
&= \varphi_{g\tau-1}^F + \sum_{i \in I_{g\tau/\tau-1}^F} \left(\widehat{\beta}_g Z_{ig\tau}\right) \varphi_{ig\tau-1}.
\end{aligned} \tag{C.3}
$$

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 $Q_{g\tau}^F \equiv \sum_{i \in I_{g\tau/\tau-1}^F} Z_{ig\tau} \varphi_{ig\tau-1}$.

C.4 Step 4

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

$$
(s_\tau^F - \widehat{s_{\tau-1}^F})^{\text{Within}} = \sum_g s_{g\tau-1}(\widehat{\varphi}_{g\tau}^F - \varphi_{g\tau-1}^F), \tag{C.4}
$$

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

$$
(\widehat{s_t^F - s_k^F})^{\text{Within}} = \sum_{\tau=k}^{t} \sum_g s_{g\tau-1}(\widehat{\varphi}_{g\tau}^F - \varphi_{g\tau-1}^F), \tag{C.5}
$$

where $k = t - 3$.

C.5 Aggregate Variance

We are interested in calculate the variance of (C.5):

$$
\text{Var}\left\{\sum_{\tau=k}^{t} \sum_g s_{g\tau-1}\widehat{\varphi}_{g\tau}^F\right\} = \sum_{\tau=k}^{t} \text{Var}\left\{X_\tau\right\} + 2\sum_{p\neq\tau}^{t} \sum_{\tau=k}^{t} \text{Cov}\left\{X_\tau, X_p\right\}, \tag{C.6}
$$

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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

$$
\begin{aligned}
\mathrm{Var}\left\{\sum_{\tau=k}^{t}\sum_{g} s_{g\tau-1}\widehat{\varphi}_{g\tau}^F\right\} &= \sum_{\tau=k}^{t}\sum_{n}\sum_{g} s_{g\tau-1}^2 \left(Q_{g\tau}^F\right)^2 \mathrm{Cov}\left\{\widehat{\beta}_g, \widehat{\beta}_n\right\} \\
&\quad + 2\sum_{p\neq\tau}\sum_{\tau=k}^{t}\sum_{n}\sum_{g} s_{g\tau-1}s_{gp-1}Q_{g\tau}^F Q_{np}^F \mathrm{Cov}\left\{\widehat{\beta}_g, \widehat{\beta}_n\right\} \qquad (\text{C.7}) \\
&= \sum_{p=k}^{t}\sum_{\tau=k}^{t}\sum_{n}\sum_{g} s_{g\tau-1}s_{gp-1}Q_{g\tau}^F Q_{np}^F \mathrm{Cov}\left\{\widehat{\beta}_g, \widehat{\beta}_n\right\}
\end{aligned}
$$

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Table 1. Aggregate Sales and Product Summary Statistics

	(1)	(2)
I. All Products		
	Total	Sales
Year	Sales	Growth
2006	2.23E+08	–
2007	2.80E+08	0.2297
2008	3.18E+08	0.1259
2009	2.80E+08	-0.1257
2010	2.82E+08	0.0075
II. Domestic Products		
	Total	Sales
Year	Sales	Growth
2006	1.37E+08	–
2007	1.71E+08	0.2247
2008	1.98E+08	0.1494
2009	1.83E+08	-0.0822
2010	1.85E+08	0.0132
III. Foreign Products		
	Total	Sales
Year	Sales	Growth
2006	8.62E+07	–
2007	1.09E+08	0.2375
2008	1.19E+08	0.0881
2009	9.75E+07	-0.2024
2010	9.72E+07	-0.0033

Notes: This table presents summary statistics for all products aggregated across all types of stores at an annual level, where a year is defined from June-May (e.g., June06–May07) in order to maximize coverage. Column (1) presents total sales in Euros. Column (2) presents the annual growth rate of sales.

Table 2. Product Group Summary Statistics

Code	Name	Share	Foreign Share	Code	Name	Share	Foreign Share
10	Meat, fresh and frozen	0.010	0.01	37	Pet foods	0.013	0.86
11	Fish	0.020	0.12	38	Pet accessories	0.002	0.85
12	Processed meat	0.043	0.03	40	Dry ingredients	0.006	0.68
13	Prepared food	0.011	0.03	41	Seasoning & preserve	0.046	0.43
14	Fresh bread	0.077	0.02	42	Sweets	0.047	0.66
21	Dairy products	0.085	0.02	43	Snacks	0.009	0.45
20	Eggs and eggs preparations	0.020	0.00	44	Dried fruit and nuts	0.009	0.18
22	Yogurts & dairy snacks	0.049	0.12	45	Natural & pharm. prods.	0.002	0.76
23	Edible fats	0.016	0.18	48	Brewery + mild alc. bevs.	0.053	0.16
24	Cheese	0.046	0.15	49	Alcoholic products	0.150	0.65
25	Frozen foods	0.018	0.40	50	Soft drinks	0.037	0.47
26	Ice cream	0.014	0.08	60	Tissues	0.013	0.73
30	Grain products	0.026	0.36	62	Disposable tableware, etc.	0.007	0.71
31	Biscuits and wafers	0.016	0.17	63	Intimate hygiene	0.007	0.98
32	Canned (jarred) foods	0.023	0.34	64	Body wash and care	0.027	0.98
33	Juices	0.023	0.21	65	Cosmetics	0.006	0.89
34	Hot drinks	0.044	0.86	66	Jewelry & optical prods.	0.002	0.81
35	Baby foods and drinks	0.009	1.00	68	Detergents	0.001	0.56
36	Baby care products	0.015	0.91		**Aggregate:**	**1.000**	**0.37**

Notes: This table presents summary statistics for two-digit product groups, aggregated across stores over the sample period May 2006–May 2011. The 'Share' column presents a product group's share of total sales over the sample period. The 'Foreign Share' column presents the share of foreign sales within a product group over the sample period. The 'Aggregate' foreign share is a 'Share'-weighted average of product groups' foreign shares.

Table 3. CES and Non-Homothetic Models' Regression Results

	CES Model	NH Model	
	(1)	(2)	(3)
	$\Delta \ln(p_{igt}/P_{gt})$	$\Delta \ln(p_{igt}/P_{gt})$	$\ln \bar{p}_{ig} \times \Delta \ln C_t$
10th pctile	-3.595	-3.588	-3.547
25th pctile	-2.814	-2.837	-0.798
50th pctile	-1.925	-1.955	0.968
75th pctile	-1.115	-1.102	3.266
90th pctile	-0.067	-0.092	6.478
Observations	236,595	236,595	
Group×time pairs	7,294	7,294	
Groups	384	384	
R^2	0.099	0.103	
χ^2_{384}: 4931.30, Prob $> \chi^2 = 0.0000$			

Notes: This table presents summary statistics on the distribution of estimated coefficients of the regression model (19). Column (1) presents the price coefficients for the CES model, while Columns (2) and (3) present the price and income coefficients, respectively, for the non-homothetic model. All specifications are run with product group×time fixed effects. The χ^2 statistic tests the null of no difference in coefficients being systematic across the two models, which is rejected. In Column (1), 270 coefficient are significant at the 10% level or below. In Columns (2) and (3), the number of coefficients that are significant at the 10% level or below are 273 and 101, respectively. Standard errors are calculated by clustering at the group×time level.

Table 4. CES and Non-Homothetic Models' Regression Results for Significnat Coefficients

|	CES Model	NH Model	
	(1)	(2)	(3)
	$\Delta \ln(p_{igt}/P_{gt})$	$\Delta \ln(p_{igt}/P_{gt})$	$\ln \bar{p}_{ig} \times \Delta \ln C_t$
10th pctile	-3.873	-3.802	-5.897
25th pctile	-3.124	-3.099	1.725
50th pctile	-2.435	-2.416	3.347
75th pctile	-1.788	-1.778	5.490
90th pctile	-1.369	-1.334	11.340
Groups	270	273	101

Notes: This table presents summary statistics on the distribution of estimated coefficients of the regression model (19), for the significant-only coefficients reported in Table 3. Column (1) presents the price coefficients for the CES model, while Columns (2) and (3) present the price and income coefficients, respectively, for the non-homothetic model. All specifications are run with product group×time fixed effects. Standard errors are calculated by clustering at the group×time level.

Table 5. CES and Non-Homothetic Models' Regression Results: Robustness Checks

	NH Non-linear Model			CES Model	NH Model	
	(1)	(2)	(3)	(4)	(5)	(6)
	$\Delta \ln(p_{igt}/$ $P_{gt})$	$\ln \bar{p}_{ig} \times$ $\Delta \ln C_t$	$\ln \bar{p}_{ig} \times$ $(\Delta \ln C_t)^2$	$\Delta \ln(p_{igt}/$ $P_{gt})$	$\Delta \ln(p_{igt}/$ $P_{gt})$	$\ln \bar{p}_{ig} \times$ $\Delta \ln C_t$
10th pctile	-3.595	-4.626	-70.502	-3.645	-3.670	-3.859
25th pctile	-2.821	-1.229	-28.508	-2.894	-2.888	-0.975
50th pctile	-1.934	0.957	0.000	-1.927	-1.973	0.899
75th pctile	-1.131	3.605	29.658	-1.115	-1.097	3.129
90th pctile	-0.088	7.125	81.780	-0.008	0.000	6.331
Observations		236,595			236,595	
GT F.E.		7,218			−	
GT-HF F.E.		−			11,419	
R^2		0.106		0.102	0.105	

Notes: This table presents summary statistics on the distribution of estimated coefficients of the regression model (19) with additional controls. Columns (1)-(3) presents results include a squared income term (Column 3), and use product group×time fixed effects ('GT F.E.'). Columns (4)-(6) present the baseline specifications for the CES and NH models with product group×time×domestic/foreign fixed effects ('GT-HF F.E.'). Standard errors are calculated by clustering at the group×time level in Columns (1)-(3), and by product group×time×domestic/foreign level in Columns (4)-(6).

Figure 1. Food and Beverages CPI and Aggregate Price Index from Scanner Data for Latvia
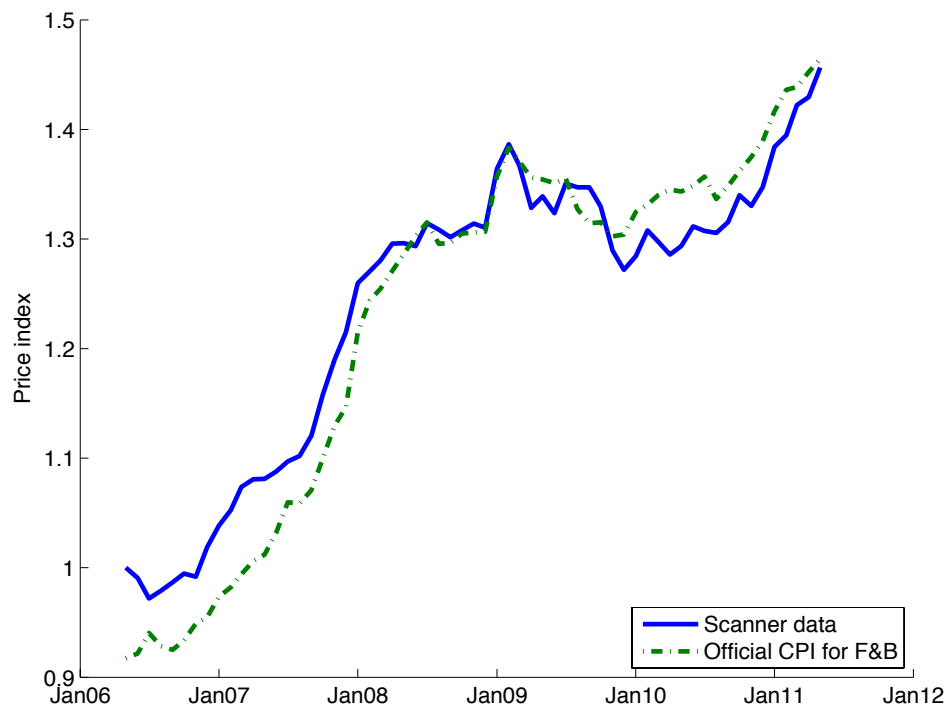

Notes: This figure plots the Latvian aggregate CPI for F&B, and an aggregate price index constructed using the scanner data. Sources: Central Statistical Bureau of Latvia and authors' calculations.

Figure 2. Food and Beverages Imports: Customs and Scanner Data
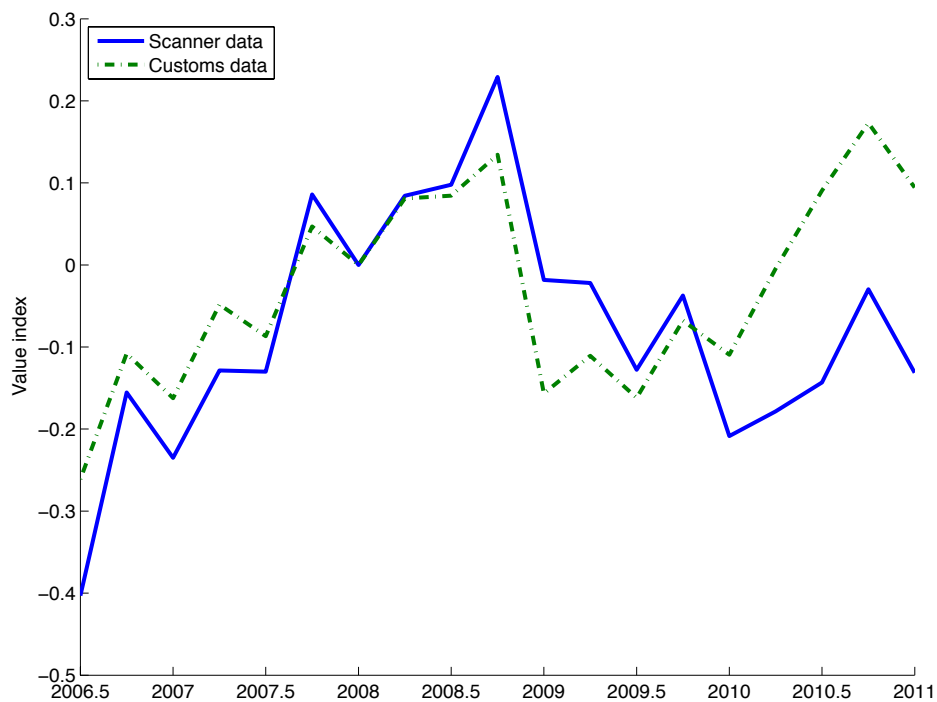
Notes: This figure plots value indexes of (i) aggregate imports of F&B for final goods (based on UN BEC classification), and (ii) expenditures on foreign goods in the scanner data. Note that both series are scaled such that 2008Q1 value is zero. Sources: Global Trade Information Services (http://www.gtis.com), UN Broad Economic Classification (http://unstats.un.org/unsd/cr/registry/regcst.asp?Cl=10), and authors' calculations.

Figure 3. Food and Beverages Fall During the Crisis
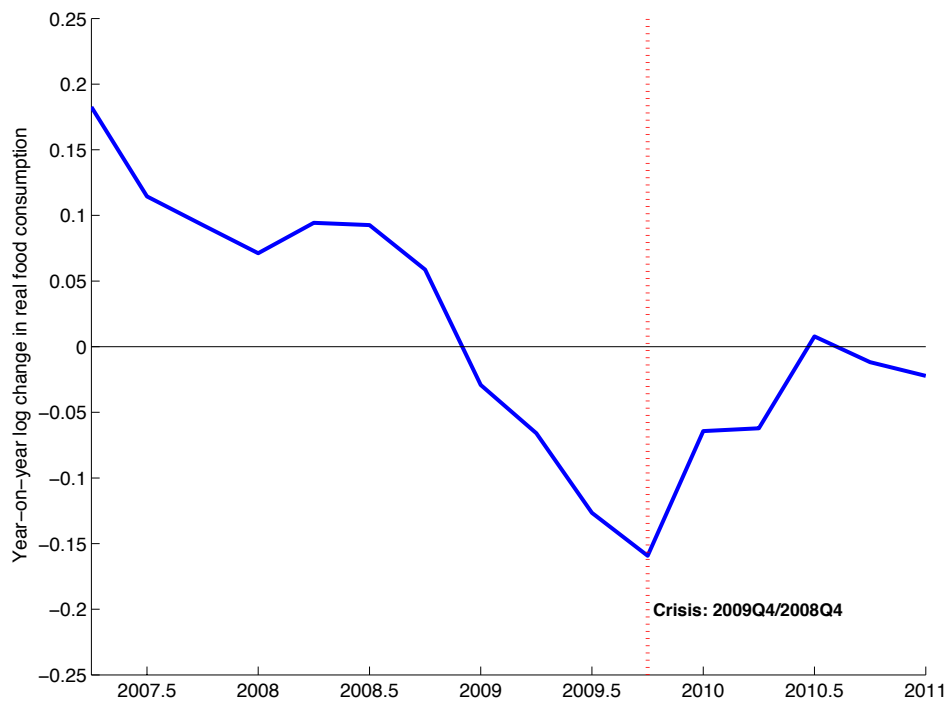
Notes: This figure plots the year-on-year log change of total real F&B expenditures over the whole sample as measured using the scanner data. Source: authors' calculations.

**Figure 4.** Expenditure Switching: Total, Within and Across Product Groups
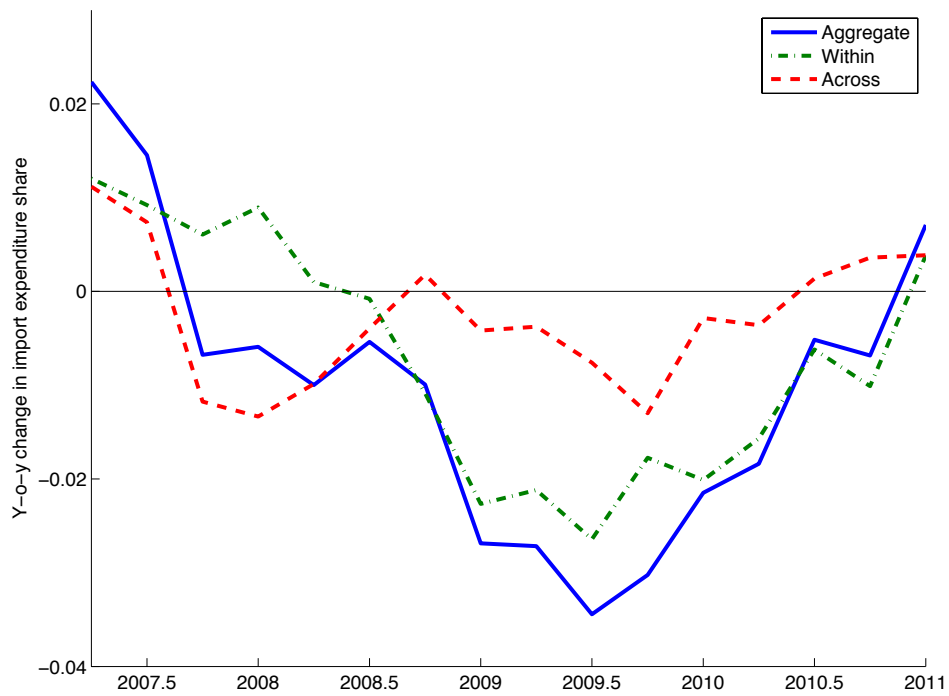
Notes: This figure plots the year-on-year change of the import share of total F&B expenditures over the whole sample as measured using the scanner data. The total change in the import share is broken into the contribution due to switching expenditures 'across' product groups and 'within' product groups (i.e., by substituting between goods), calculated using (1).

Figure 5. Relative Price Change: Total, Within and Across Product Groups

Notes: This figure plots the year-on-year change of the relative price of foreign goods for F&B expenditures over the whole sample as measured using the scanner data. The total change in the relative price is broken into the contribution due to changes 'across' product groups and 'within' product groups (i.e., by substituting between goods), calculated using (5).

Figure 6. Distribution of Within Product Group Interquartile Range Unit Values at the 4-Digit Product Code Level
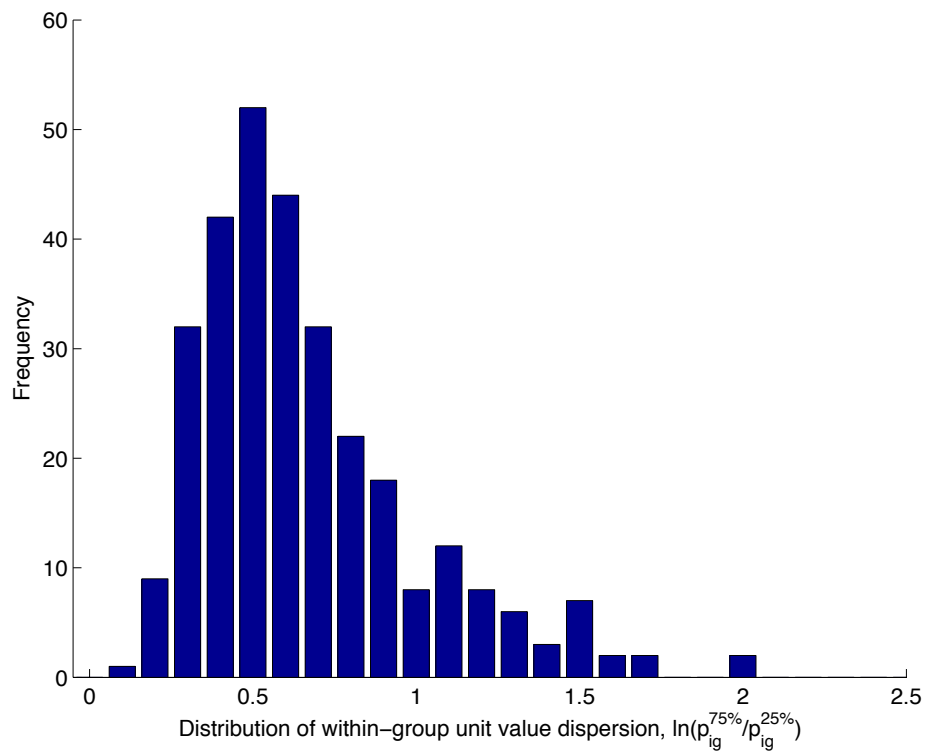

Notes: This figure plots the distribution of within product interquartile range unit values across all 4-digit product groups over the entire sample. Interquartile range of a given product group is defined as the difference between unit value of the items at the 75th and 25th percentiles of the product group, averaged over 20 quarters.

Figure 7. Distribution of Within Product Group Relative Unit Values of Imported and Domestic Items at the 4-Digit Product Code Level
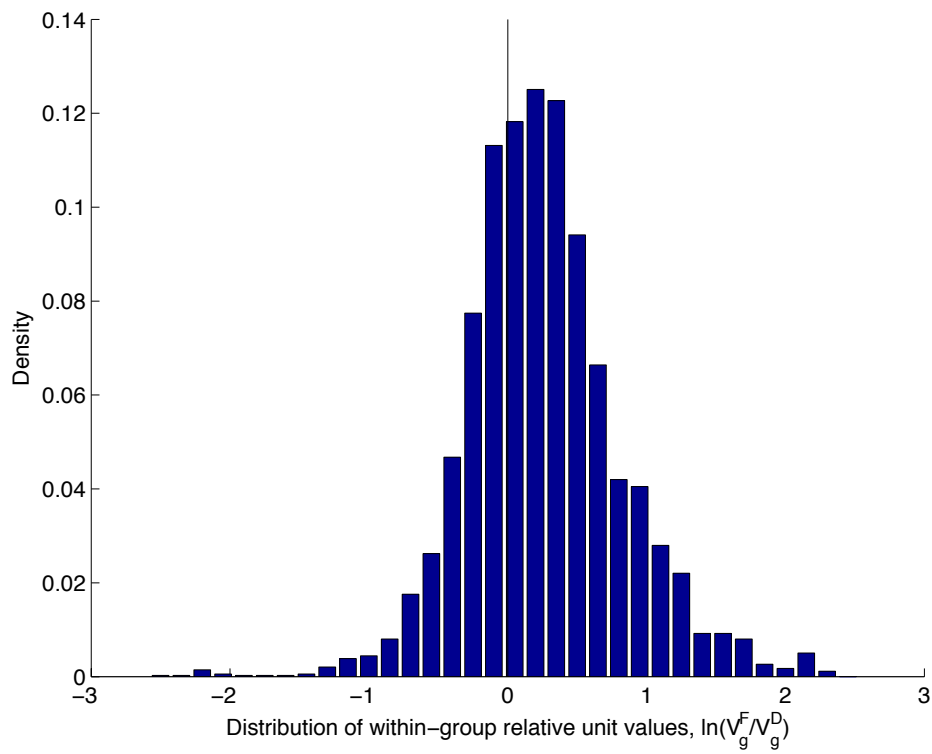

Notes: This figure plots the distribution of the relative unit values of foreign (V_g^F) and domestic (V_g^D) items at the four-digit product code level. The histogram is constructed using average relative unit values for each product group over the whole sample period.

Figure 8. Deviations in Price Changes, as Measured by Product Group's Unit Value and Price Index
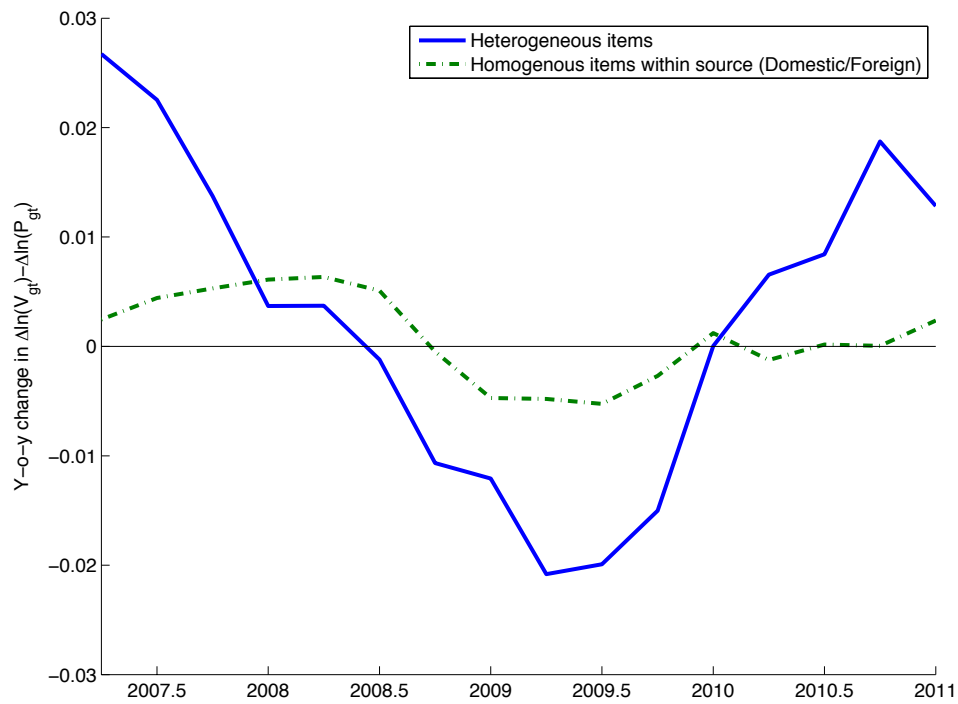

Notes: This figure compares y-o-y deviations in within-group price changes, as summarized by a group's unit value and price index. Results aggregated over product groups by weighting each group with its share in aggregate food expenditures. A sub-component of aggregate deviations, capturing deviations due to item source, is derived by assuming that items within a source (i.e., domestically produced or imported) are homogeneous.

Figure 9. Model Estimated and Actual Within Components of Expenditure Switching
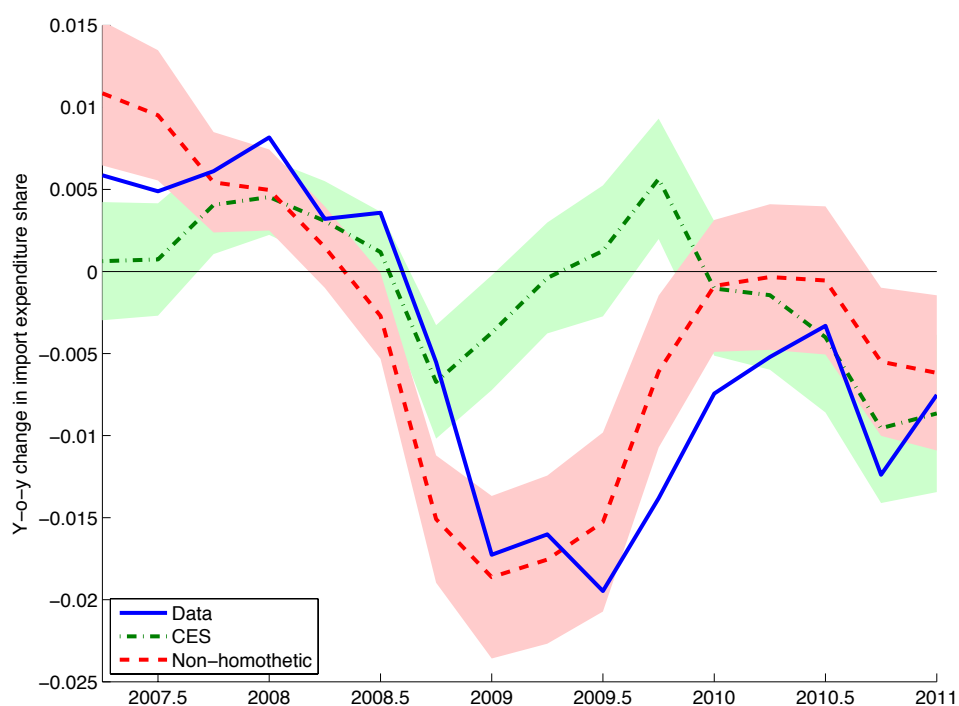
Notes: This figure plots the *within* component of expenditure switching observed in the data and estimated using the model based on (19), for the CES and Non-homothetic models. The shaded areas are two standard error bands, calculated analytically based on clustered standard errors at the group×time level.

Figure 10. Non-Homothetic Model's Within Components of Expenditure Switching: Income and Price Effects
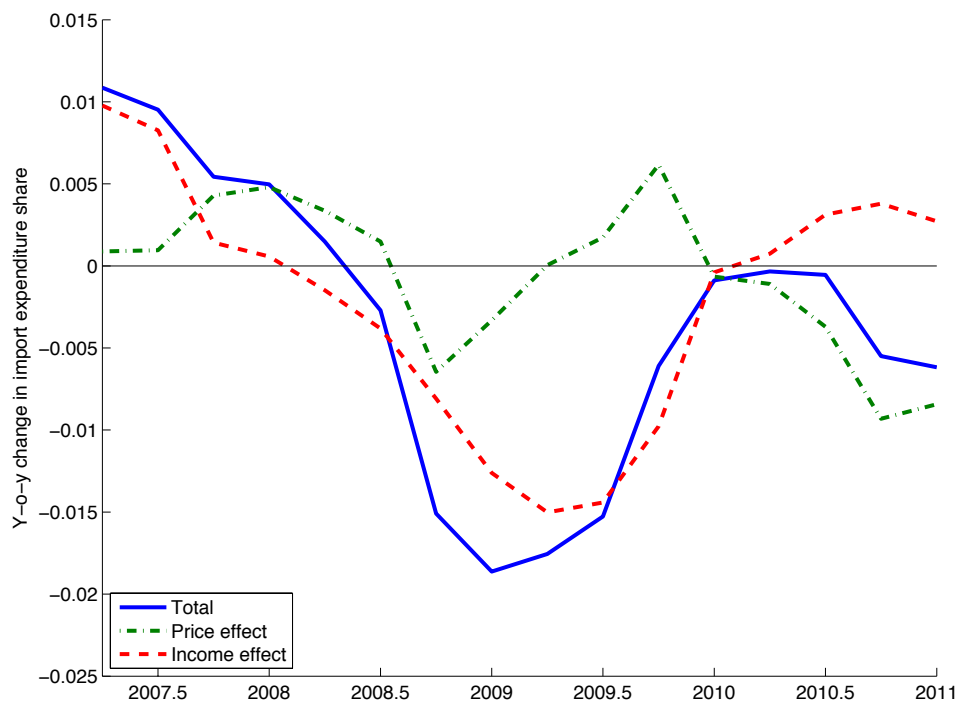
Notes: This figure plots the estimated *within* component of expenditure switching predcited by the Non-homothetic model, breaking it down into contributions due to (i) a price effect, and ((ii) an income effect. The model is estimated using the full model (19).

Table A1. Product-Group-Level Regression Results

	(1)	(2)	(3)	(4)
$\Delta \ln(P_{gt}/P_t)$	-1.089**	-1.078**	-1.268**	-1.259**
	(0.197)	(0.170)	(0.160)	(0.167)
Observations	7,294	7,294	7,294	7,294
Group F.E.			384	384
Time F.E.		19		19
R^2	0.030	0.029	0.040	0.052

Notes: This table presents results of estimating the regression model $\Delta \ln s_{gt} = \alpha_t + \alpha_g + \Delta \ln(P_{gt}/P_t) + \varepsilon_{gt}$, at the product group level. Column (1) presents OLS results; Column (2) includes time fixed effects; Column (3) includes product group fixed effects, and Column (4) – our preferred specification – includes time and group fixed effects. Standard errors are calculated by clustering at the group level (except for Column (3), which are clustered at the time level). A ** indicates significance at the 1% level.

Table A2. CES and Non-Homothetic Models' Regression Results: OLS and Different Configurations of Fixed Effects

| | I. CES Model | | | | II. NH Model | | | | | | | |
| | (1) | (2) | (3) | (4) | (1) | | (2) | | (3) | | (4) | |
	β_{1g}	β_{1g}	β_{1g}	β_{1g}	β_{1g}	β_{2g}	β_{1g}	β_{2g}	β_{1g}	β_{2g}	β_{1g}	β_{2g}
5th pctile	-4.060	-4.056	-4.021	-4.009	-4.152	-6.210	-4.150	-6.074	-4.100	-6.030	-4.095	-6.096
10th pctile	-3.524	-3.536	-3.517	-3.515	-3.543	-3.573	-3.569	-3.400	-3.530	-3.445	-3.508	-3.468
25th pctile	-2.815	-2.821	-2.798	-2.804	-2.834	-0.706	-2.815	-0.533	-2.808	-0.712	-2.804	-0.605
50th pctile	-1.897	-1.910	-1.912	-1.926	-1.917	0.904	-1.923	1.127	-1.922	0.886	-1.938	1.130
75th pctile	-1.127	-1.117	-1.112	-1.116	-1.130	2.789	-1.126	2.959	-1.100	2.861	-1.102	2.973
90th pctile	-0.052	-0.049	-0.035	-0.038	-0.093	5.696	-0.073	6.063	-0.047	5.699	-0.075	6.055
95th pctile	0.352	0.329	0.428	0.431	0.369	9.801	0.344	9.916	0.462	10.056	0.444	10.141
Observations	236,595				236,595							
Group F.E.	–	–	384	384	–		–		384		384	
Time F.E.	–	19	–	19	–		19		–		19	
R^2	0.097	0.097	0.098	0.099	0.102		0.097		0.103		0.104	

Notes: This table presents summary statistics on the distribution of estimated coefficients of the regression model (19), which is estimated using group×time effects in Table 3, for different arrays of fixed effects. Panel I presents results for the CES model estimation, while Panel II presents corresponding results for the non-homothetic model. Column (1) presents regression with no fixed effects; Column (2) presents results with only time effects; Column (3) presents results with only group effects; Column (4) presents results with (uninteracted) group and time effects. Standard errors are calculated by clustering at the group×time level.
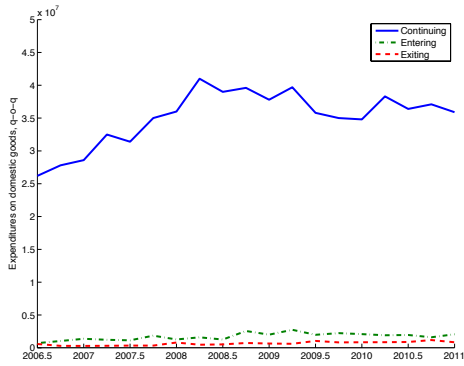
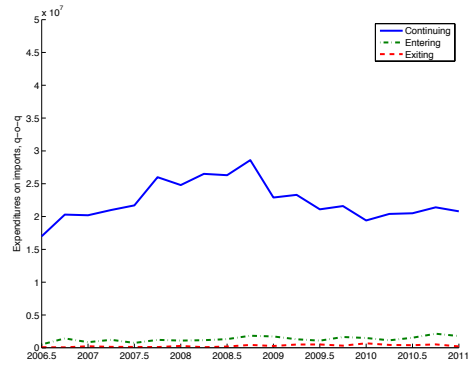**Figure A1.** Domestic and Import Goods: Continuing, Entry, and Exit
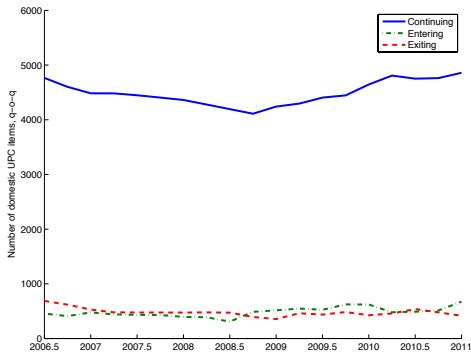
**(a)** Domestic Count
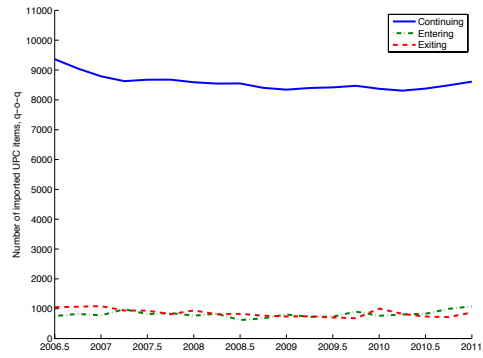
(b) Foreign Count

(c) Domestic Expenditure
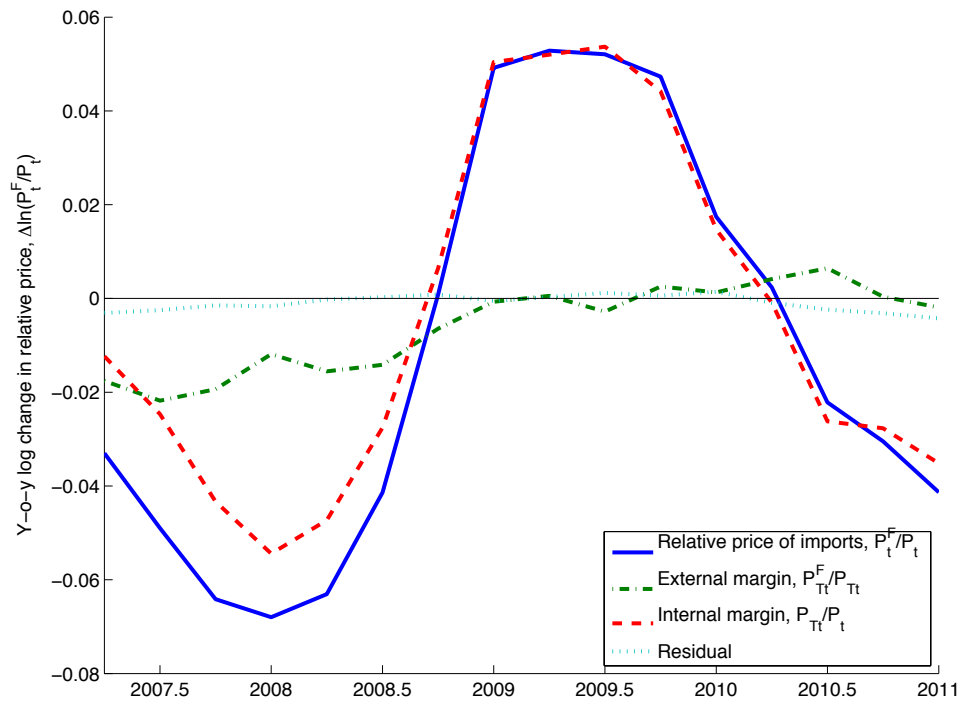
(d) Foreign Expenditure

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.

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Figure A2. Growth Rate of Expenditure Switching: Total and Intensive and Extensive Margins
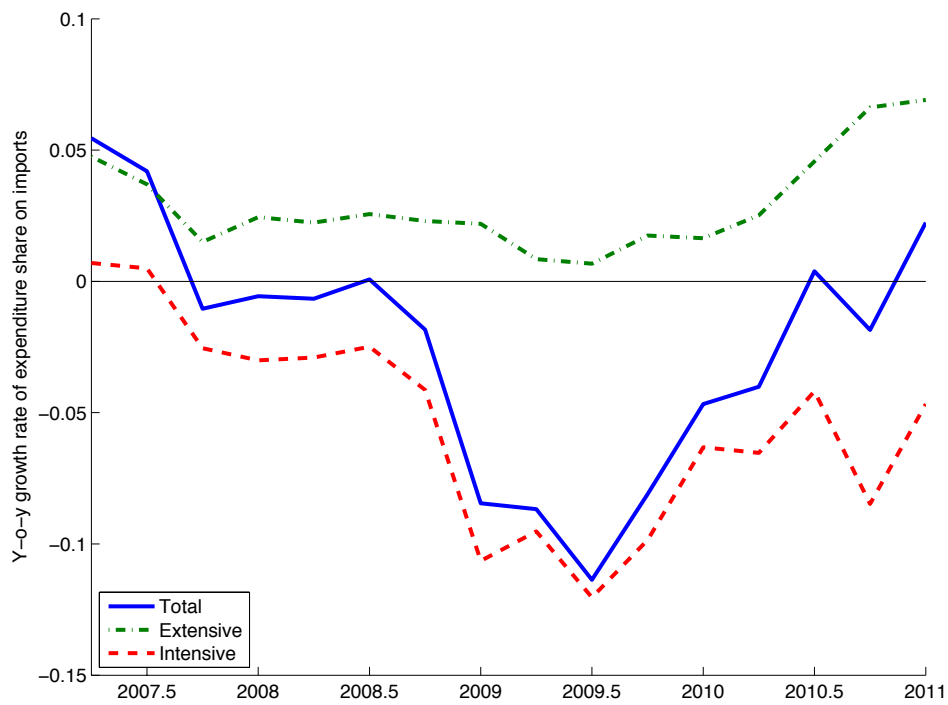
Notes: This figure plots the growth rate of the total expenditure share on imported goods, as well as the contribution to growth due to changes for continuing goods – the 'intensive margin' – and due to net entry and exit of goods – the 'extensive margin'. Growth rates are calculated using quarterly data and are then accumulated over a four-quarter overlapping rolling window.

Figure A3. Relative Price Changes and Tradability

Notes: This figure plots the y-o-y change of the relative price of foreign goods for F&B expenditures over the whole sample as measured using the scanner data. The total change in the relative price is broken into the contribution due to the 'internal margin' or changes in the relative price of tradables to nontradables and the 'external margin' or changes in the relative price of domestic and imported tradables. Nontradables are defined as 4-digit product groups with a shelf-life of <180 days.