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Income-Induced Expenditure Switching†

By Rudolfs Bems and Julian di Giovanni*

This paper shows that an income effect can drive expenditure switching between domestic and imported goods. We use a unique Latvian scanner-level dataset, covering the 2008–2009 crisis, to document several empirical findings. First, expenditure switching accounted for one-third of the fall in imports, and took place within narrowly defined product groups. Second, there was no corresponding within-group change in relative prices. Third, consumers substituted from expensive imports to cheaper domestic alternatives. These findings motivate us to estimate a model of nonhomothetic consumer demand, which explains two-thirds of the observed expenditure switching. Estimated switching is driven by income, not changes in relative prices. (*JEL* E21, F14, F31, F32, I11, L81)

The exchange rate plays a central role in discussions of external balance adjustments across countries. When prices are 缓慢 sluggish to adjust, a currency depreciation provides a potentially fast way to reduce domestic prices, relative to foreign ones, which in turn increases demand for domestic goods at home and abroad and leads to expenditure switching. Therefore, a common policy 法规 prescription for countries with fixed exchange rate systems, and facing balance of payments crises, has been to devalue their exchange rates in order to facilitate external adjustment. The international macroeconomics theory underlying this conventional external adjustment channel assumes that a change in a country's income affects the consumption of

*Bems: Graduate School of Public Policy, University of Tokyo, 7-3-1 Hongo, Bunkyo-ku, Tokyo 113-0033 Japan (e-mail: rudolfs.bems@gmail.com); di Giovanni: Department of Economics and Business, Universitat Pompeu Fabra Ramon Trias Fargas 25-27, Barcelona 08005, Spain, ICREA, Barcelona GSE, CREI, and CEPR (e-mail: julian.digiovanni@upf.edu). We are grateful to Rimi Baltic, Anders Alexanderson, Vyacheslav Dombrovsky, and Anders Paalzow for helping us acquire the scanner-level data. For helpful comments we would like to thank two anonymous referees, Cristina Arellano, Olivier Blanchard, Paula Bustos, Oli Coibion, Fabrizio Coricelli, Giancarlo Corsetti, Doireann Fitzgerald, Jessie Handbury, Jonathan Heathcote, Rob Johnson, Andrei Levchenko, Philippe Martin, Akito Matsumoto, Isabelle Mejean, Maurice Obstfeld, Franck Portier, Steven Phillips, Jay Shambaugh, Mick Silver, Alan Taylor, Jaume Ventura, Kei-Mu Yi, and seminar participants at the 2013 AEA Meetings, Bank of England, Bank of Latvia, CREI, ECB conference on "Heterogeneity in Currency Areas and Macroeconomic Policies," ESSIM 2014, European Commission conference on "Current Account Imbalances and International Financial Integration," FRB of Governors, IMF, Minneapolis Fed, 2013 NBER Summer Institute IFM meetings, Philadelphia Fed International Trade Workshop, University of Tokyo, Swiss National Bank, and SNB-CEPR conference on "Exchange Rates and External Adjustment." Jair Rodriguez and Marola Castillo provided superb research assistance. This work was partly carried out while Bems was visiting the Bank of Latvia and the Stockholm School of Economics in Riga. Hospitality of both institutions is gratefully acknowledged. All remaining errors are our own. The authors declare that they have no relevant or material financial interests that relate to the research described in this paper.

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domestic and foreign goods proportionally, so that a relative price change is the only source of expenditure switching (Engel 2003; Obstfeld and Rogoff 2007).

This paper revisits the relationship between relative prices, income changes, and expenditure switching during a balance of payments crisis. We exploit a unique item-level dataset to demonstrate that *income-induced* expenditure switching is needed to understand the data, and that a model with nonhomothetic preferences better matches the observed expenditure switching than a constant elasticity of substitution (CES) model, which is typically used in international macroeconomics.

We focus on the 2008–2009 balance of payment crisis in Latvia, during which the country defied the conventional policy prescription and maintained its exchange rate 与欧元挂钩 pegged to the euro.[1] To the surprise of many economists, within two years from the outset of the crisis, a 20 percent net trade-to-gross domestic product (GDP) deficit was reduced to balanced trade and GDP growth resumed (see Figure 1). The bulk of the external adjustment took place on the import side, as the share of imports in GDP declined from 65 percent in 2007 to 45 percent in 2009.[2] The adjustment in aggregate relative prices was 缓和的 subdued, as Latvia's real exchange rate remained broadly unchanged over the period 2008–2011.[3] Latvia's experience has generated recent interest because it is one of the few examples where a large external adjustment was achieved faster than expected and without a nominal devaluation or a significant adjustment in relative prices, thus potentially shedding light on a successful adjustment process in a monetary union, and particularly in the southern periphery countries of the eurozone who also faced an appreciated real exchange rate and price rigidity.

This paper zooms in to the microeconomic level to better understand what drove the adjustment in imports. First, we quantify how much expenditure switching took place between domestic and imported goods. Second, we explore the margins—across or within product groups—at which expenditure switching and relative price changes took place. Finally, we ask whether the observed relative price changes can explain the observed expenditure switching through the lens of the conventional theory and whether the income effect had a role to play.

We measure relative price and consumption changes across goods using a scanner-level dataset on food, beverages, and other supermarket items, covering the 2006:II–2011:I period. These data provide both prices and quantities at the individual item level and crucially identify the country origin of each item and detailed product groups to which items belong to. The key advantage of the dataset is that it allows for measurement of expenditure switching and relative prices with internally consistent data, which we find to be representative of aggregate expenditure and price movements in the food sector. It is also important to note that a similar exercise using available trade and macroeconomic data is not possible, since such data do not

[1] See Blanchard, Griffiths, and Gruss (2013) for a forensic account of Latvia's boom, bust, and recovery over the period 2000–2013.

[2] The crucial role of import compression in driving the adjustment was also observed in the other Baltic states, as well as the eurozone periphery countries (Kang and Shambaugh 2014).

[3] As panel C of Figure 1 demonstrates, the main driver of Latvia's real exchange rates over the period was the nominal effective exchange rate, which in fact *appreciated* by 5 percent during 2008–2009 because of *third-country* exchange rate movements (i.e., the euro was appreciating viz. Latvia's non-euro trade partners). Meanwhile, wages in manufacturing were cut by 7 percent initially, but quickly recovered to precrisis levels (Blanchard, Griffiths, and Gruss 2013).

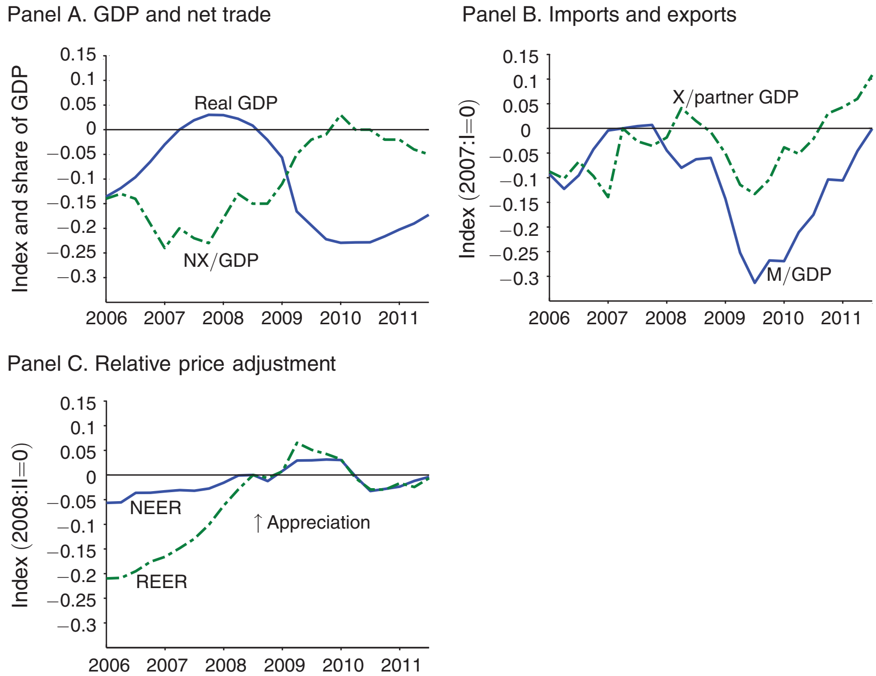

FIGURE 1. LATVIA'S BALANCE OF PAYMENTS CRISIS OF 2008–2009

Notes: This figure plots the evolution of the key macro variables for Latvia around the 2008–2009 crisis episode. NX/GDP denotes net trade, as a share of GDP. In panel A, Real GDP is a seasonally adjusted log volume index. In panel B, M/GDP denotes imports of goods and services as a share of GDP, and X/partner GDP denotes exports of goods and services as a share of trade partner GDP. All variables are normalized to zero in 2007:I, and are based on data from the Central Statistical Bureau of Latvia, except partner GDP, which is from the IMF's International Financial Statistics (IFS). In panel C, NEER denotes the nominal effective exchange rate index, and REER denotes the CPI-based real effective exchange rate index. Both indexes are normalized to zero in 2008:II, and an increase in the indexes represents an appreciation. The NEER and REER data are from the IMF's IFS.

identify quantities and prices of comparable domestic and imported goods in final consumption.

We find that, during the crisis period, expenditures on imports fell by 26 percent, while overall expenditures on food contracted by 18 percent, so that expenditure switching from imported to domestic goods accounted for one-third of the total fall in imports in the scanner dataset.[4] We then use the item-level dimension of the data to present two main findings. First, the majority of the expenditure switching was driven by substitution between goods *within* narrowly defined product groups. Second, there was no corresponding change in the relative price of imports within product groups. The change in the relative price of imports—a modest 3.8 percent rise—was driven almost entirely by changes in prices *across* product groups.

The observed expenditure switching within product groups without a corresponding relative price adjustment presents a puzzle. Why did consumers switch to domestic substitute items within product groups, if such items did not become less expensive than their imported counterparts? Our proposed answer focuses on the

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

shift in consumed item mix within product groups, as summarized by a third empirical finding: *consumers substituted from expensive imported items to cheaper domestic alternatives*. We find that within narrowly defined product groups, imported items are on average 28 percent more expensive than comparable domestic items, and that consumers substituted toward cheaper similar items, irrespective of origin, during the crisis. This substitution generated expenditure switching from imported to domestic items without any adjustment in relative prices.

Motivated by the empirical findings, we set up a demand-side model of the economy to formally quantify contributions of relative prices and the substitution toward cheaper goods to the observed expenditure switching, and to link consumers' substitution behavior to the observed fall in aggregate income. We model an expenditure allocation problem of a representative consumer. Given item-level prices and the crisis-induced fall in income, a consumer decides how to allocate expenditures across and within product groups. The consumer's choice depends on relative prices as well as an income-driven demand for quality. With quality considerations switched off, the model is a conventional CES demand system.

We estimate the model parameters in a panel regression setting using disaggregated item-level data, and control for a host of potential factors that might otherwise bias the estimation results, as well as running instrumental variable regressions on subsets of the data. We then use the estimated parameters to construct a predicted aggregate measure of expenditure switching over the sample period. The results are striking. First, the conventional CES model performs poorly: though the estimated price elasticities are similar to available estimates in the literature, and are statistically significant, the model's predicted expenditure switching does not match the switching observed in the data, particularly during the crisis episode. Second, the nonhomothetic model is better able to match observed expenditure switching during the crisis—it captures two-thirds of what is observed in the data. We further find that the income channel, not relative prices, drives expenditure switching in the nonhomothetic model.

Our findings are consistent with the *flight from quality* hypothesis put forth in Burstein, Eichenbaum, and Rebelo (2005), who present facts on consumer shopping patterns during the 2001 Argentinean crisis. They argue that ignoring consumers' substitution toward lower quality (thus cheaper) goods introduces a bias in measured CPI, and relative price movements. Besides complementing Burstein, Eichenbaum, and Rebelo's (2005) findings, we are able to shed further light on the impact of this *flight* by identifying a switch from foreign to domestic items and linking the flight to changes in income. The absence of a large devaluation, combined with the drastic fall in income, makes the Latvian crisis episode an ideal case for identifying the income-induced channel of expenditure switching.

Although the results in this paper are based on scanner-level data for a particular sector for one country, they may speak to broader issues in international macroeconomics. First, the international trade literature has documented that poorer countries tend to be net importers of higher quality goods across all sectors of the economy (see, e.g., Hummels and Klenow 2005; Hallak 2006; Feenstra and Romalis 2014). Therefore, for these countries, the income effect may play an important role in external adjustment, regardless of exchange rate movements. Given the recent crisis in the eurozone, there has indeed been an adjustment in the periphery countries who were subject to large negative income shocks, and the income-induced expenditure switching is a possible channel to help explain this phenomenon. Second, there is also a large literature documenting the Alchian and Allen (1964) effect, which posits that traded goods tend to be higher quality than domestic ones: commonly referred to as "shipping the good apples out" (see Hummels and Skiba 2004 for some recent empirical evidence). If this is indeed the case, then income-induced expenditure switching may be a more general phenomenon affecting external adjustment across countries. Furthermore, the proposed channel may have asymmetric effects across countries given different income levels. For example, a rich country like the United States may in fact be a net importer of *lower* quality goods, so expenditure switching may go in the opposite direction than what we find for Latvia. Third, in support of the flight from quality hypothesis, some recent work has pointed to a fall in the quality composition of large EU countries' exports during the Great Trade Collapse (GTC).[5] Such work, however, remains silent about implications for expenditure switching, as there is no matching data for domestic goods.

Our work can be related to several strands of existing literature. We contribute to the large literature in international macroeconomics on external adjustment. This literature is comprised of an extensive list of theoretical studies on expenditure switching and the role of exchange rate policy (see Engel 2003 for a review; and Burstein, Eichenbaum, and Rebelo 2007; Kehoe and Ruhl 2009; Mendoza 2005; and Obstfeld and Rogoff 2007 for work studying sudden stop episodes). Previous work has also suggested that besides relative prices, the extensive margin can have an impact on the external adjustment (e.g., see Krugman 1989; Corsetti, Martin, and Pesenti 2013). The nonhomothetic channel we introduce in our work is distinct from this mechanism. Furthermore, we show empirically that our results are robust to extensive margin considerations.

This paper also contributes to the extensive literature that measures the adjustment in relative prices (see Burstein and Gopinath 2014 for a recent survey).[6] Our contribution to this literature is twofold. First, we provide direct empirical evidence for expenditure switching, which can be linked to a relative price adjustment. To the best of our knowledge, such evidence at the microeconomic level is nonexistent.[7] Second, we contribute to the scarce literature on external adjustment under a fixed exchange rate regime, which makes our study particularly relevant for the current policy debate on the external adjustment process in a currency union.[8]

[5] See, for example, Berthou and Emlinger (2010) and Esposito and Vicarelli (2011). Though, focusing on US import data, Levchenko, Lewis, and Tesar (2011) reject the hypothesis that the fall in imports was skewed toward higher quality goods. Meanwhile, in the context of an emerging market, Chen and Juvenal (2015) use detailed microdata on wine exports from Argentina and measures of wine quality to show that there was a fall in the quality of exports during the GTC.

[6] Of particular note, work by Parsley and Popper (2006) studies relative price movements under a fixed exchange rate regime (Hong Kong), and recent papers by Berka, Devereux, and Engel (2012) and Cavallo, Neiman, and Rigobon (2014) contrast real exchange rate adjustments in and outside of the eurozone.

[7] 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 Mejean (2015) and Feenstra et al. (2014) 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).

[8] See Farhi, Gopinath, and Itskhoki (2014) for recent work studying alternative ways of generating devaluations in the absence of exchange rate flexibility.

Our findings also emphasize the relevance of nonhomothetic preferences in macroeconomics and complement a rapidly growing literature on trade and income inequality that models nonhomotheticities in the demand for quality.[9] We are not the first to study macroeconomic implications of nonhomotheticities in consumption. For example, nonhomothetic preferences are widely used in the literature on growth and structural change.[10] We instead focus on crises (i.e., large economic fluctuations). Diaz Alejandro (1965) is an early study of how income effects can affect external rebalancing.[11] He investigates how consumption behavior differences between wage and nonwage earners affect the demand of different sectors' imports. In the absence of data on income heterogeneity, we instead focus on the large change in aggregate income and explore its implications for demand of high/low priced items in narrowly defined product groups.

The rest of the paper is organized as follows. Section I describes the data. Section II presents the paper's three main empirical findings about expenditure switching and the accompanying price adjustment. Section III employs an estimated demand model to more formally quantify the contribution of relative price changes and income effects in explaining the observed expenditure switching. Section IV concludes.

I. 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 (universal product code) items sold by one of Latvia's largest retailers, Rimi.[12] The data are collected across three types of stores that the retailer owns and runs: a chain of hypermarkets (H), a chain of typical supermarkets (S), and a chain of discounter (D) stores.[13] We treat items sold at each store as distinct. Data for each type of store are aggregated across the respective type's sales-per-item across the country, so there is no geographical distinction by type of store. The provided data cover the six-year period May 2006 through May 2011, which spans Latvia's boom, bust, and the beginning of its subsequent recovery. The coverage of goods is primarily for food and beverages (F&B), but the dataset also contains other consumer goods typically available at supermarkets, such as toiletries. Besides quantity and price information, the dataset provides information on the unit of measure (e.g., kilograms), the net content of each UPC item, and a short item description with an accompanying retailer assigned *material* code.

The retailer also provides two-, three-, and four-digit classifications of the items into product groups. An example of a two-digit product group would be "hot drinks," which at the three-digit level is further broken down into "tea," "coffee," and "cacao." The three-digit group "tea" is further broken down at the four-digit

[9] For example, Hallak (2006); Fajgelbaum, Grossman, and Helpman (2011); Fajgelbaum and Khandelwal (2016); Faber (2014).

[10] See Herrendorf, Rogerson, and Valentinyi (2014) for a recent survey.

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

[12] Rimi Baltic is a major retail operator in the Baltic states based in Riga, Latvia. It is a subsidiary of Swedish group ICA. Rimi Baltic operates 235 (as of year 2013) retail stores in Estonia (83 stores), Latvia (113 stores), and Lithuania (39 stores), and has distribution centers in each country.

[13] The S and H stores carry a wider variety of goods than D stores, and the same 13-digit UPC item can vary in price across the three types of stores.

level into types of tea. For example, there is "unflavored black tea," "flavored black tea," "herbal tea," "fruit tea," etc. Since this is not a widely used product classification, it is useful to contrast it to the standard combined nomenclature (CN8) classification of trade flows. We find that for a majority of product groups, the four-digit supermarket product classification is comparable to or even narrower than the CN8 trade flow classification at the most detailed eight-digit level. Thus, four-digit supermarket scanner product groups are based on very narrow product definitions.

The 13-digit 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 barcode 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 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.[14]

An alternative approach to identifying the item's origin suggests that the UPC is a valid proxy. We focus on the domestic/imported origin for a subset of four-digit product groups that explicitly group items by origin (e.g., imported and domestic beer). Such product groups account for 11.6 percent of total F&B expenditures in our sample, 6.2 percent of which are identified as domestic, and 5.4 percent as foreign. We find that for product groups that are identified as domestic, 97.3 percent of expenditures carry local UPCs. For product groups that are identified as imported, 97.2 percent of expenditures carry foreign UPCs. This suggests that for a small market, such as Latvia, the UPCs can correctly identify the domestic/foreign origin for more than 97 percent of expenditures. We also note that, despite the significantly reduced sample size, the main findings of this paper remain unchanged if this alternative identification of origin is used.

A. *Data Cleaning and Consolidation*

As with any large micro dataset, data cleaning is needed. First, given our focus on expenditure switching between domestic and imported goods, we are forced to drop all store products for which we cannot identify their origin. Such items are produced/labeled by the retailer, with the bulk falling into specific product groups such as "store bake," "fruits and berries," and "vegetables and root crops." The 13-digit UPC identifies such items as store products, but provides no information about the origin of ingredients. Store products and product groups dominated by such items account for 29 percent of total food expenditures in the dataset over the whole sample period.[15]

Second, we drop items (i) without a UPC and (ii) with either quantity or price less than or equal to zero. Imposing these two conditions left total sales virtually unchanged, decreasing them by 0.3 percent.

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

[15] We reintroduce store products back into the sample when examining results that do not require information about items' origin, such as shifts in within-group item mix (see Section IIC).

We next consolidate the scanner-level data for homogeneous items. We start by consolidating data by the triplet of (i) UPC, indexed with i; (ii) store type (s); and (iii) time period (t), because in the original dataset 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 recomputing item unit values from the 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 percent 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 consolidation:

Product code	Product group description	"Material" code	Item description	Net content	UPC	Quantities 2009/4	Quantities 2009/5	Quantities 2009/6	Average price 2009/4	Average price 2009/5	Average price 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	1,331	640		0.47	0.47	
2101	Fat-free milk	211961	MILK VALMIERA 0.5% 1L	1 L	4750074005062		994	2,152		0.47	0.47

The 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 different 13-digit UPCs. For the purpose of this paper, such items are 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 UPC-store items in our sample by 14 percent.[16]

B. *Summary Statistics*

Table 1 presents annual data for all products in the resulting dataset, 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 2006 through April 2007. Looking at columns 1 and 2, we see that the value of sales increased until 2008–2009 when the crisis hit, and there is then a pick up in 2010–2011 as the Latvian economy began to recover. The same pattern holds for both domestic and foreign sales. Column 3 reveals an elevated price increase during the precrisis boom, driven by prices of domestic goods. Once the crisis hits, there is deflation, again driven by the domestic component of food expenditures. Column 4 reports the number of unique UPC-store items sold by the retailer each year, as well as their domestic/imported breakdown. Overall, there

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

Table 1—Aggregate Sales and Product Summary Statistics

	Total sales (1)	Sales growth (2)	Price change (3)	UPC-store items (4)
Panel A. All products				
2006	235.7	—	—	42,049
2007	296.3	0.257	0.131	39,423
2008	336.7	0.137	0.079	38,781
2009	297.0	−0.118	−0.016	36,794
2010	302.0	0.017	0.024	36,689
Panel B. Domestic products				
2006	135.7	—	—	12,945
2007	169.7	0.251	0.172	12,308
2008	197.4	0.163	0.078	11,868
2009	182.6	−0.075	−0.039	12,220
2010	184.0	0.008	0.039	12,830
Panel C. Foreign products				
2006	100.0	—	—	29,104
2007	126.6	0.266	0.077	27,115
2008	139.4	0.101	0.080	26,913
2009	114.4	−0.179	0.018	24,574
2010	118.0	0.032	0.001	23,859

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 May to April (e.g., May 2006 through April 2007) in order to maximize coverage. Column 1 presents total sales in millions of euros; Column 2 presents the annual growth rate of sales; Column 3 presents the annual inflation rate based on our constructed price indexes from store items; and Column 4 presents the total count of unique UPC-store items appearing in the sample in a given year.

are 80,890 unique UPC-store items sold over the five-year sample period, of which 26,244 are of domestic origin and the remaining 54,646 are imported.

Next, Table 2 presents summary statistics for two-digit product groups over the whole sample period. The Share column reports the share of each product group's sales viz. total sales over the period. Alcoholic products make up the largest share of total sales, accounting for 14.7 percent of aggregate revenue. The Foreign Share column measures the foreign content of a given product group. Though the food and beverage sector is generally considered a tradable sector (Berka and Devereux 2013; Crucini, Telmer, and Zachariadis 2005), we find considerable heterogeneity in import intensity among product groups at this relatively high level of aggregation. Foreign contents range from a low of 2 percent (dairy products) to a high of 100 percent (baby foods). While the foreign share of total sales is 40 percent, imports account for a mere 7 percent of the least import-intensive quarter of food sales.

C. *Aggregated Scanner-Level Data and Macroeconomic Trends*

Food and beverages account for approximately 30 percent of total household expenditures in Latvia,[17] therefore the scanner-level data cover an important

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

Table 2—Product Group Summary Statistics

Code	Name	Share	Foreign share	Code	Name	Share	Foreign share
10	Meat, fresh and frozen	0.010	0.07	37	Pet foods	0.015	0.88
11	Fish	0.022	0.26	38	Pet accessories	0.002	0.85
12	Processed meat	0.047	0.18	40	Dry ingredients	0.006	0.68
13	Prepared food	0.011	0.05	41	Seasoning and preservation	0.044	0.44
14	Fresh bread	0.074	0.04	42	Sweets	0.047	0.67
21	Dairy products	0.019	0.02	43	Snacks	0.009	0.52
20	Eggs and eggs preparations	0.084	0.06	44	Dried fruit and nuts	0.009	0.22
22	Yogurt and dairy snacks	0.051	0.20	45	Natural and pharm. prods.	0.002	0.77
23	Edible fats	0.016	0.24	48	Brewery + mild alc. bevs.	0.053	0.21
24	Cheese	0.045	0.27	49	Alcoholic products	0.147	0.66
25	Frozen foods	0.020	0.50	50	Soft drinks	0.039	0.52
26	Ice cream	0.016	0.24	60	Tissues	0.013	0.74
30	Grain products	0.027	0.41	62	Disposable tableware, etc.	0.007	0.72
31	Biscuits and wafers	0.016	0.22	63	Intimate hygiene	0.006	0.98
32	Canned (jarred) foods	0.022	0.36	64	Body wash and care	0.025	0.98
33	Juices	0.022	0.23	65	Cosmetics	0.006	0.89
34	Hot drinks	0.042	0.86	66	Jewelry and optical prods.	0.001	0.82
35	Baby foods and drinks	0.009	1.00	68	Detergents	0.001	0.62
36	Nappies and baby care prods.	0.014	0.91		Aggregate	1.000	0.40

Notes: This table presents summary statistics for two-digit product groups, aggregated across stores over the sample period May 2006 through 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.

component of total consumption. Furthermore, given the size of the retailer, the scanner-level dataset directly adds up to 15 percent 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-level data.

First, as Figure 2 shows, the constructed aggregate price index closely mimics the official F&B's consumer price index (CPI).[18] Second, retail market share data, kindly provided to us by IGD Retail Analysis, show that during the period 2007–2011, the retailer maintained a stable grocery retail market share of around 20 percent (see Table 3). Further, the number of stores operated by the retailer did not vary systematically with the crisis.

Next, Figure 3 plots the total revenue of foreign products across all stores and aggregate F&B imports used for final consumption. The two series are highly correlated, and the scanner-level data pick up the large fall in imports over the crisis period.[19]

Given that a sizable part in the fall of trade during the Great Trade Collapse was due to durables, we also examine the breakdown of Latvia's annual imports between nondurable and durable goods during the boom-bust-recovery period in Table 4. Imports of food and beverages (excluding tobacco) constitute 25 percent of

[18] The monthly CPI from scanner-level data is constructed using the multilateral GEKS price index. For an in-depth discussion of this index, see Ivancic, Diewert, and Fox (2011).

[19] 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 (Alessandria, Kaboski, and Midrigan 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.

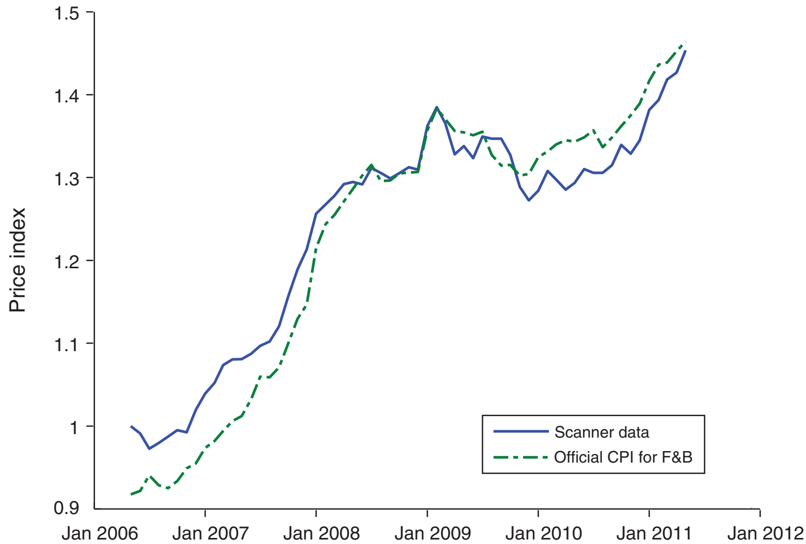

Figure 2. Food and Beverages CPI and Aggregate Price Index from Scanner-Level Data for Latvia

Note: This figure plots the Latvian aggregate CPI for F&B, and an aggregate price index constructed using the scanner-level data.

Sources: Central Statistical Bureau of Latvia and authors' calculations

Table 3—Grocery Retail Market Share of Rimi in Latvia

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

Source: IGD Retail Analysis

nondurable imports, which in turn make up roughly one-half of total imports, and accounted for a substantial portion of total import growth over the sample period. The fall in durable imports was indeed larger than for nondurables during the crisis in 2009. However, the fall in nondurable imports was substantial (i.e., 28 percent), and accounted for roughly one-third of the total fall in imports, which is consistent with what happened in the rest of the world (see Bems, Johnson, and Yi 2013). In comparison, imports of F&B fell by 21 percent in 2009 (see Figure 3), which was somewhat less than the fall in nondurables, but still of a similar magnitude. Therefore, although our scanner-level dataset covers only one sector of the economy, it is an important one, both for total private consumption and aggregate imports.

Finally, why use scanner-level data rather than more commonly available macroeconomic data to study expenditure switching? The micro-level data allow us to measure prices and expenditures on domestic and imported goods at a very disaggregate level within a single large dataset using consistent final consumer prices. In contrast, macroeconomic data would require combining data on trade flows with household expenditure data, which creates multiple issues. The main issue is that household expenditures are measured in final consumer prices and bundle together

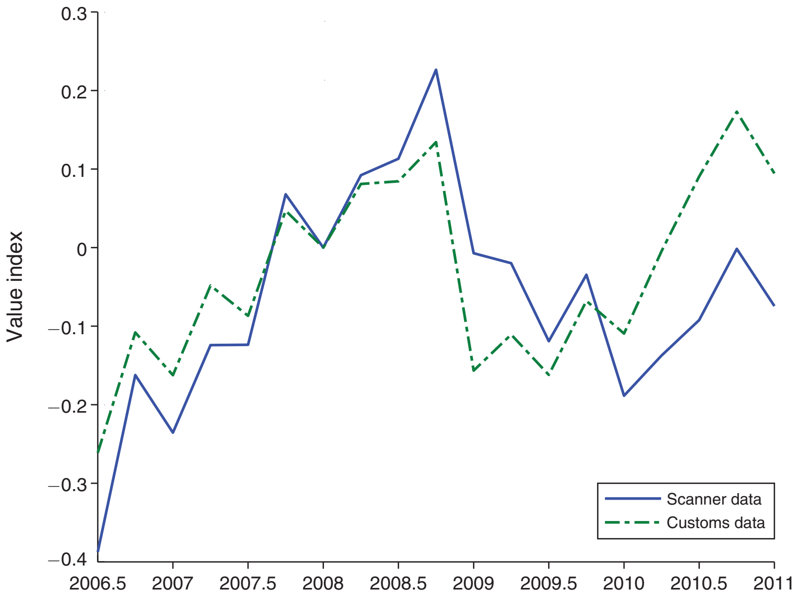

Figure 3. Food and Beverages Imports: Customs and Scanner-Level Data

Notes: This figure plots value indexes of (i) aggregate imports of F&B for final goods (based on UN Broad Economic Classification (BEC)), and (ii) expenditures on foreign goods in the scanner-level data. Note that both series are scaled such that the 2008:I value is zero.

Sources: Global Trade Information Services (http://www.gtis.com), UN BEC (http://unstats.un.org/unsd/cr/registry/regcst.asp?Cl=10), and authors' calculations

Table 4—Latvian Imports: Durable and Nondurable Goods

	Growth rate			Share	
	Total	Nondurable	Durable	Nondurable	Durable
2006	0.27	0.20	0.34	0.44	0.56
2007	0.19	0.15	0.22	0.42	0.58
2008	−0.01	0.12	−0.12	0.48	0.52
2009	−0.44	−0.28	−0.62	0.56	0.44
2010	0.21	0.14	0.28	0.53	0.47
2011	0.30	0.25	0.36	0.50	0.50
2012	0.14	0.15	0.12	0.51	0.49

Notes: This table presents the breakdown of Latvian imports between durable and nondurable goods. Durability is defined following Engel and Wang (2011) and Lechenko, Lewis, and Tesar (2011). The first three columns present the growth rate of total, nondurable, and durable imports, respectively. The last two columns present the share of nondurable and durable goods in total imports.

Source: Eurostat and authors' calculations

expenditures on imports with a large domestic retail margin, while trade flows are measured at the dock (Berger et al. 2009; Burstein, Eichenbaum, and Rebelo 2005). Furthermore, National Income Account data estimate expenditures indirectly (and at a relatively high level of aggregation) using infrequent surveys. Another issue is that inventories can drive a wedge between final expenditures and trade flows, especially during sudden stop episodes (Alessandria, Kaboski, and Midrigan 2010).

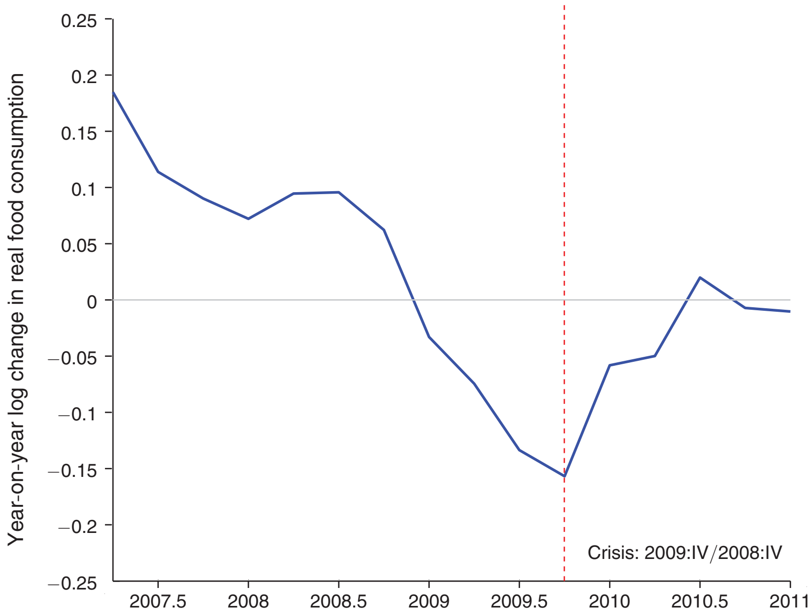

Figure 4. Food and Beverages Fall during the Crisis

Note: This figure plots the y-o-y log change of total real F&B expenditures over the whole sample as measured using the scanner-level data.

II. Empirical Findings

The Latvian economy experienced a sharp contraction during the sudden stop, felt across all sectors of the economy, including consumption of food and beverages. Figure 4 uses quarterly data to plot the year-on-year (y-o-y) log change in real aggregate food consumption in the scanner-level data.[20] The figure depicts a classic *boom-bust* episode. Consumption was growing before the crisis, at which point it experienced a substantial drop, bottoming out at −16 percent in real terms over the period 2008:IV–2009:IV. To eliminate seasonality and provide a consistent time frame for the results reported in this paper, we label these four consecutive quarters with the largest cumulative fall in food consumption as the crisis year. Note that this is done purely for presentational convenience, as none of our main findings hinge on a specific definition of the crisis period.

The scanner-level 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.

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

The three empirical findings are:

Finding 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.

Finding 2: The expenditure switching was accompanied by a 3.8 percent 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.

Finding 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 adjustment in relative prices.

A. *Finding 1: Expenditure Switching*

We first examine the role of expenditure switching in the total fall of imports during the crisis. A contraction in imports can be linked to either an across-the-board compression in food consumption, which affects domestic and imported food proportionally, or a reallocation of food expenditures from imports toward domestic products, i.e., expenditure switching. In the scanner-level data over the 2008:IV–2009:IV period, imports fell by 26 percent, while total food expenditures fell by 18 percent. Therefore, expenditure switching accounted for 8 percentage points, or one-third, of the fall in imports.

An alternative way of quantifying the size of the expenditure switching is to consider a y-o-y change in the aggregate import expenditure share in quarterly data. To construct the import expenditure share from the item-level data, it is useful to introduce some notation. Define a product group $g \in \{1,\ldots,G\}$, 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 we can express the expenditure share of a product group as $s_{gt} = \sum_{i \in I_{gt}} s_{igt}$, and the aggregate expenditure share on imports as $s_t^F = \sum_g \sum_{i \in I_{gt}^F} s_{igt}$, where F refers to the subset of imported items.

The solid line in Figure 5 plots the y-o-y change in the aggregate import share, $s_t^F - s_{t-4}^F$, over the sample period. At the trough (i.e., 2009:IV relative to 2008:IV), 3.7 cents for every euro spent on F&B were reallocated from imports toward domestically produced food. Since the aggregate import expenditure share in the dataset is 0.40 (see Table 2), this amounts to a 9.3 percent fall in the import expenditure share.

Although there is entry and exit of items in the scanner-level data, we find that the adjustment at the *intensive* margin accounts for the bulk of the expenditure switching in Latvia (see online 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

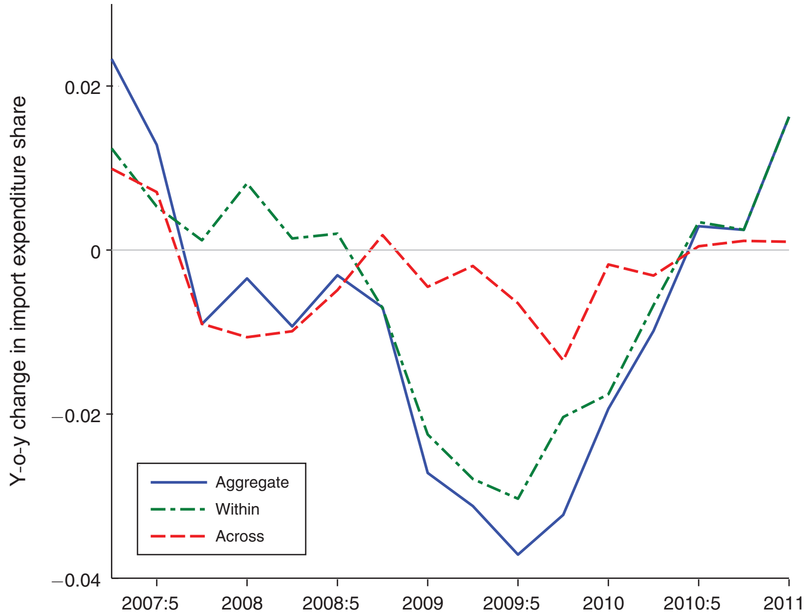

FIGURE 5. EXPENDITURE SWITCHING: TOTAL, WITHIN AND ACROSS PRODUCT GROUPS

Notes: This figure plots the y-o-y change of the import share of total F&B expenditures over the whole sample as measured using the scanner-level 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).

extensive margin played a small role (see Bems, Johnson, and Yi 2013 for a recent review).

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 to expenditure switching indirectly 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.

Formally, define the share of imports within a product group as $\varphi_{gt}^F \equiv s_{gt}^F / s_{gt}$. 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 additive components of interest—expenditure switching within and across product groups—as follows:

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

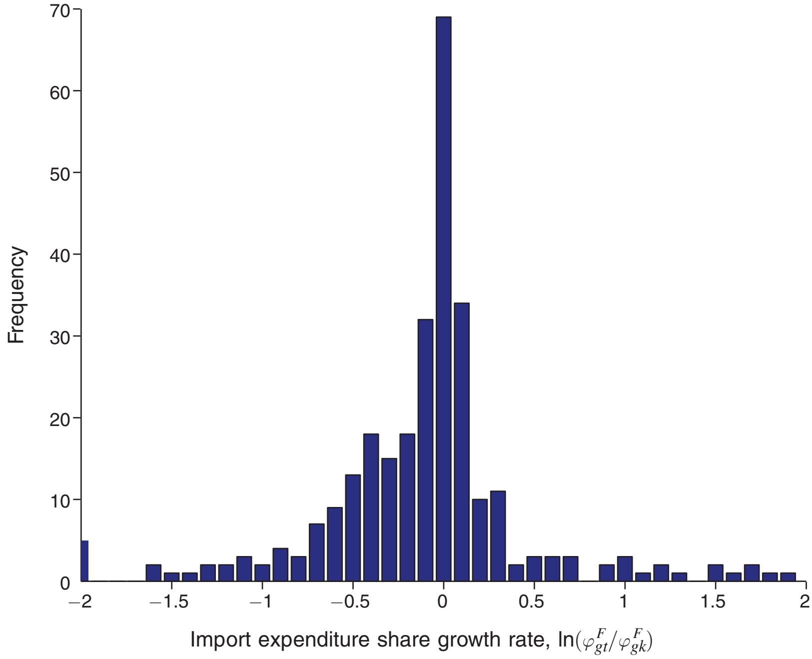

FIGURE 6. DISTRIBUTION OF IMPORT EXPENDITURE SHARE GROWTH RATES FOR FOUR-DIGIT PRODUCT GROUPS OVER THE CRISIS PERIOD

Notes: This figure summarizes the y-o-y expenditure switching growth rates at the four-digit product group level over the crisis period, i.e., $t = 2009$:IV and $k = 2008$:IV. There are 291 four-digit product groups with nonzero expenditures on domestic and imported goods. The growth rate in the median group is −0.04 and the mean is −0.15.

Figure 5 plots this decomposition for y-o-y changes in s^F_t, where a product group g is defined at the narrowest four-digit level. We find that the bulk of expenditure switching took place within the narrow sectors (dashed-dotted 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 zoom in on this key expenditure switching decomposition result, Figure 6 shows the distribution of import share growth rates during the crisis period, $\varphi^F_{g,\,09:\mathrm{IV}}/\varphi^F_{g,\,08:\mathrm{IV}}$, for the 291 four-digit product groups that make up Finding 1. The histogram reveals sizable dispersion in the import share growth rates, with the share of imports contracting in two-thirds of product groups. The growth rate for the median groups is −4 percent and the mean growth rate is −15 percent.

Given the structure of the scanner data, we can further investigate whether the switching within product groups is taking place predominantly across store types (e.g., buying the same UPC item at a Discounter store instead of a Supermarket),[21] or across items within a given store type. A simple extension of (1), detailed in online Appendix B, allows us to quantify contributions from these two substitution margins to the overall expenditure switching. Results show that more than

[21] See Coibion, Gorodnichenko, and Hong (2015).

90 percent of expenditure switching took place *within*, rather than *across*, store types. This finding can be explained in terms of the potential for savings that the two substitution margins offer in the scanner data. Substituting across stores can provide savings of up to 11 percent for a median item. As we will show in Section IIC, this margin of savings is a fraction of what consumers can save by substituting across items in the narrowly defined four-digit product groups.

B. *Finding 2: Relative Price Adjustment*

We next examine price movements of domestic and import goods at the aggregate and product group level. In order to do so, we 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.[22] 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, $w_{igt} \equiv 0.5(s_{igt} + s_{igt-1})$ 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 7 plots the y-o-y change in $\ln(P_t^F / P_t)$. The relative price increases by 3.8 percent y-o-y during the crisis period (2008:IV–2009:IV), and by 5.3 percent from trough to peak.[23]

[22] The findings of this paper are robust to the use of alternative price index definitions, such as the multilateral GEKS, Fisher, or Laspeyres, in the construction of aggregate prices.

[23] Note that the relative price increase in imported food contrasts with the overall appreciation of the CPI-based real exchange rate as discussed in footnote 3 above. These contrasting movements can be reconciled by the fact that food exports (i.e., imports in our scanner data) are not included in CPI in source countries. In addition, imports from non-euro countries are likely invoiced in euros, limiting the effect of exchange rate movements on import prices.

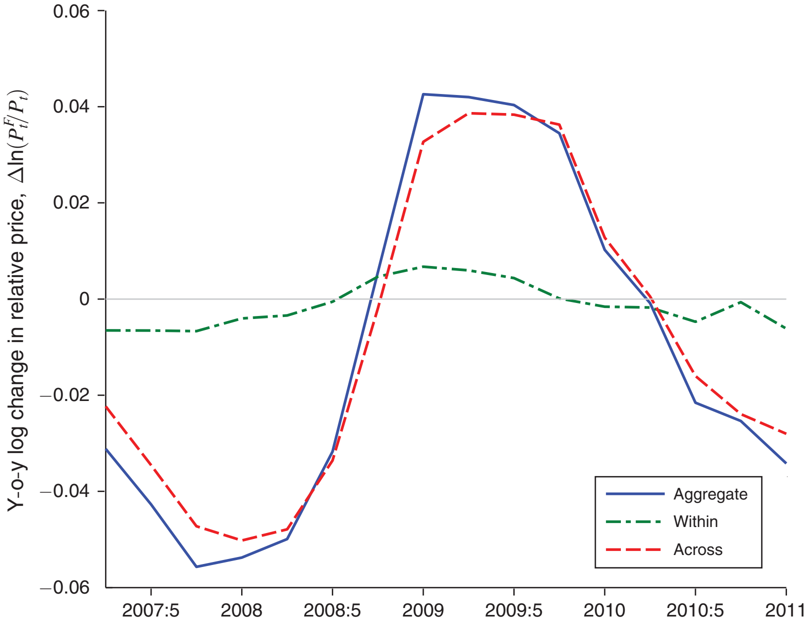

Figure 7. Relative Price Change: Total, within and across Product Groups

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-level 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).

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 $w_{gt}^F / w_t^F = \sum_{i \in I_{gt}^F} w_{igt} / \left(\sum_g \sum_{i \in I_{gt}^F} w_{igt} \right)$ 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 as

$$(5) \quad \ln \frac{P_t^F}{P_t} - \ln \frac{P_k^F}{P_k} = \sum_g \frac{w_{gt}^F}{w_t^F} \left(\ln \frac{P_{gt}^F}{P_{gt}} + \ln \frac{P_{gt}}{P_t} \right) - \sum_g \frac{w_{gk}^F}{w_k^F} \left(\ln \frac{P_{gk}^F}{P_{gk}} + \ln \frac{P_{gk}}{P_k} \right)$$

$$= \underbrace{\sum_g \frac{w_{gk}^F}{w_k^F} \left(\ln \frac{P_{gt}^F}{P_{gt}} - \ln \frac{P_{gk}^F}{P_{gk}} \right)}_{\text{Within}} + \underbrace{\sum_g \frac{w_{gk}^F}{w_k^F} \left(\ln \frac{P_{gt}}{P_t} + \ln \frac{P_{gk}}{P_k} \right)}_{\text{Across}}$$

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

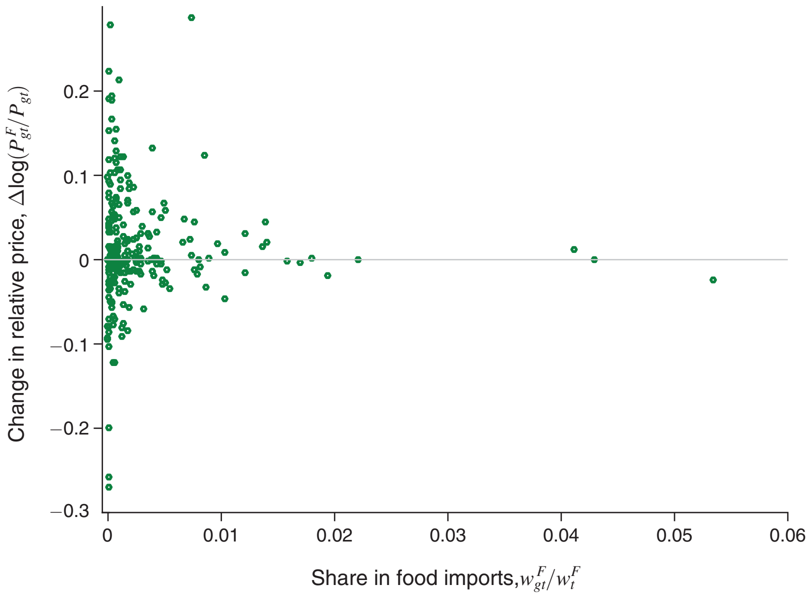

Figure 8. Import Price Adjustments within Four-Digit Product Groups during the Crisis Period

Notes: This figure reports import price adjustments for each of 265 four-digit product groups over the crisis period, i.e., 2009:IV–2008:IV. Price adjustments are plotted against each product group's share in total imports prior to the crisis, i.e., 2008:IV. The relative price of imports in the median group increases by 0.001, while the mean increase is 0.013.

In Figure 7, again using four-digit product groups, we see that the increase in the relative price of imports was almost entirely driven by price movements across product groups (dashed line). Within-product group relative prices (dashed-dotted line) did not exhibit sizable systematic deviations.[24] Our findings with regard to the price adjustment are the opposite of what occurred for expenditure shares, where switching took place within, not across product groups. From the conventional macroeconomic theory standpoint, these contrasting findings present a puzzle: why are consumers buying more domestic varieties even though domestic items are not becoming less expensive than their foreign counterparts?

To shed further light on this puzzle, Figure 8 zooms in on the two components of the within-price adjustment in (5): group-by-group changes in the relative price of imports, $\ln(P^F_{gt}/P_{gt}) - \ln(P^F_{gk}/P_{gk})$, and the corresponding product groups' shares in total imports, w^F_{gk}/w^F_k.[25] We find the absence of a systematic import price adjustment during the crisis period within four-digit product groups to be broad-based. The import price in the median group increases by a mere 0.1 percent. The mean increase is 1.4 percent, but this somewhat elevated figure is driven by price increases

[24] 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.

[25] The number of groups (265) included in Figure 8 is smaller than the number of groups (291) included in Figure 6, because for some four-digit product groups with a small number of observations, we can compute changes in expenditure shares but not the price adjustment. Excluded product groups account for 3.8 percent of total expenditures on F&B.

in small product groups. When weighted by their contribution to imports, the mean price increase during the crisis drops to 0.4 percent, as already reported in Figure 7.

We also perform several exercises to directly relate import price adjustments to expenditure switching across the four-digit product groups during the crisis. First, a simple correlation between expenditure switching growth rates, as reported in Figure 6, and import price adjustments, as reported in Figure 8, is -0.12, and becomes even weaker when small product groups are dropped from the sample. Second, we calculate contributions to expenditure switching separately for product groups with increasing and decreasing import prices. This is an instructive exercise, because regardless of elasticity assumptions, conventional macro models predict that expenditure switching should take place only in product groups where import prices increase. We find, instead, that 45 percent of the aggregate expenditure switching during the crisis took place in product groups with falling, not increasing, import prices. Overall, the data shows no sizable systematic adjustment in import prices, and no systematic relationship between import price adjustments and expenditure switching within four-digit product groups during the crisis.

C. *Finding 3: Shifts in Within-Group Item Mix*

In the absence of a systematic relative price adjustment within sectors, this section explores alternative sources of expenditure switching. The section documents a large dispersion in item unit values within four-digit product groups and a pervasive within-group substitution toward cheaper items during the crisis. This substitution is then linked to expenditure switching from more expensive imported items to cheaper domestic substitutes. In the subsequent section the proposed expenditure switching channel is estimated and quantified more formally.

Unit Value Dispersion and Flight to Cheaper Substitutes.—There are large differences in item unit values within product groups.[26] Figure 9 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 seventy-fifth and twenty-fifth percentiles of the product group's distribution of unit values. We find that for the median product group, the unit value at the seventy-fifth percentile is 53 percent larger than the twenty-fifth percentile, while the mean difference is 67 percent.

The observed dispersion in unit values is broad-based across data subsets. Table 5 summarizes the extent of the dispersion by focusing on four measures: the median and mean unit value differences between the seventy-fifth and twenty-fifth percentiles, the standard deviation of unit values within product groups, and the number of items sold in the median product group. Results are reported for all items as well as subsets of data broken down by an item's origin. Regardless of the examined subset, there are sizable differences in unit values at the seventy-fifth and twenty-fifth

[26] Examination of unit value differences within four-digit product groups requires restricting the data to comparable net units (e.g., kilograms, liters) within each product group. This is implemented by (i) dropping product groups where *pieces* are used as the measure of units and (ii) limiting each group to items measured in common units. Imposing these two restrictions decreases total expenditures in the dataset by correspondingly 7.6 percent and a further 2.1 percent.

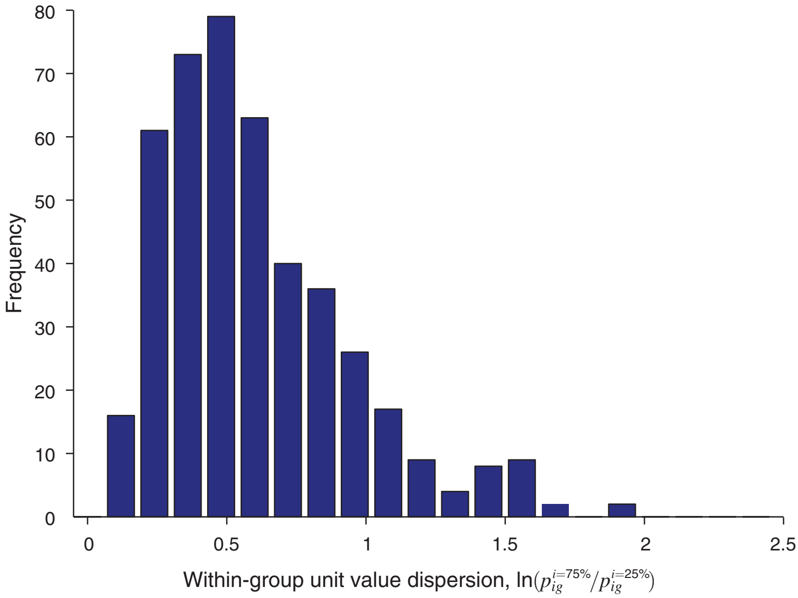

Figure 9. Distribution of Within-Product Group Interquartile Range Unit Values at the Four-Digit Product Code Level

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

Table 5—Unit Value Dispersion: All Items and Select Subsets

	Median	Mean	SD	Number of items
All items	0.53	0.67	0.58	33
By item origin				
Foreign	0.53	0.69	0.56	24
Domestic	0.39	0.50	0.37	19
Store products	0.47	0.62	0.43	18

Notes: The means and medians are based on the distribution of the within-product interquartile ranges of unit values across all four-digit product groups over the entire sample. The interquartile range of a given product group is defined as the log ratio of the unit value of the items at the seventy-fifth and twenty-fifth percentiles of the product group, as reflected in the histogram in Figure 9. The SD column measures the standard deviation of the unit values within a four-digit product group over the sample, normalized by the mean unit value of the group over the sample. The number of items column reports the number of items sold in the median product group. Store products refer to groups/items that were dropped from the baseline dataset because of missing information about origin.

percentiles, ranging between 39 percent and 53 percent for the median product group. The dispersion is some 25 percent smaller for domestic items, when compared to foreign ones.[27]

[27] A similar unit value dispersion is found if the data are split by (i) store type and (ii) product groups with higher/lower unit values of imported relative to domestic items.

The documented dispersion in unit values within narrow four-digit product groups potentially offers an alternative margin for substitution: during the crisis and in the absence of any significant price adjustment, consumers might have switched to lower unit value items. To examine this possibility, we construct a Laspeyres-type expenditure share index for group g between t and $t+1$:

$$
\begin{aligned}
(6)\quad \Delta W_{gt} &= \sum_i \left(\ln \varphi_{igt+1} - \ln \varphi_{igt}\right)\left(\ln p_{igt} - \ln \overline{p}_{gt}\right) \\
&= \underbrace{\sum_i \left(\ln \phi_{igt+1} - \ln \phi_{igt}\right)\left(\ln p_{igt} - \ln \overline{p}_{gt}\right)}_{\text{Quantity}} \\
&\quad + \underbrace{\sum_i \left(\ln p_{igt+1} - \ln p_{igt}\right)\left(\ln p_{igt} - \ln \overline{p}_{gt}\right)}_{\text{Price}} \\
&\quad + \underbrace{\ln F_g \sum_i \left(\ln p_{igt} - \ln \overline{p}_{igt}\right)}_{\approx 0},
\end{aligned}
$$

which sums changes in item expenditure shares, defined as $\varphi_{igt} \equiv p_{igt} q_{igt} / \sum_i p_{igt} q_{igt}$, multiplied by each item's unit value, relative to the group's average base year unit value, defined as $\overline{p}_{gt} \equiv \sum_i p_{igt} q_{igt} / \sum_i q_{igt}$.[28] The second line further decomposes the index into contributions from changes in items' prices (Price) and quantity shares (Quantity), defined as $\phi_{igt} \equiv q_{igt} / \sum_i q_{igt}$, and a residual term that is approximately zero by definition.[29]

This index can isolate systematic shifts in expenditure shares between high and low unit value items. When expenditures are shifting toward less expensive items, the index takes a negative value and vice versa when expenditures shift systematically toward more expensive items. The quantity and price components allow us to separate between the two key sources of shifts in expenditures. The quantity component of the index in (6) takes on negative values when quantity shares shift toward less expensive items, while the price component takes on positive values when the relative price of lower unit value items is decreasing.

We compute the quantity and price components of ΔW_{gt} for each four-digit product group g and each t using quarterly data. Indexes for individual four-digit product groups are then aggregated into an overall index using each group's weight in total F&B expenditures, and the resulting index is re-expressed in y-o-y changes to avoid seasonality. Similar results are obtained if unweighed mean or median changes in the indexes are considered instead.

The resulting aggregated quantity and price components, reported in panel A of Figure 10, show that the consumed quantities of high/low unit value items within product groups varied systematically with the boom-bust-recovery cycle. In particular, during the crisis the quantity component (solid line) turned negative, with a trough for the shift toward less expensive items in mid-2009. Panels B–D show

[28] We are indebted to a referee for suggesting this index.
[29] Since $F_g \equiv \left(\sum_i q_{igt+1} / \sum_i q_{igt}\right)\left(\sum_i p_{igt} q_{igt} / \sum_i p_{igt+1} q_{igt+1}\right)$ is a constant term.

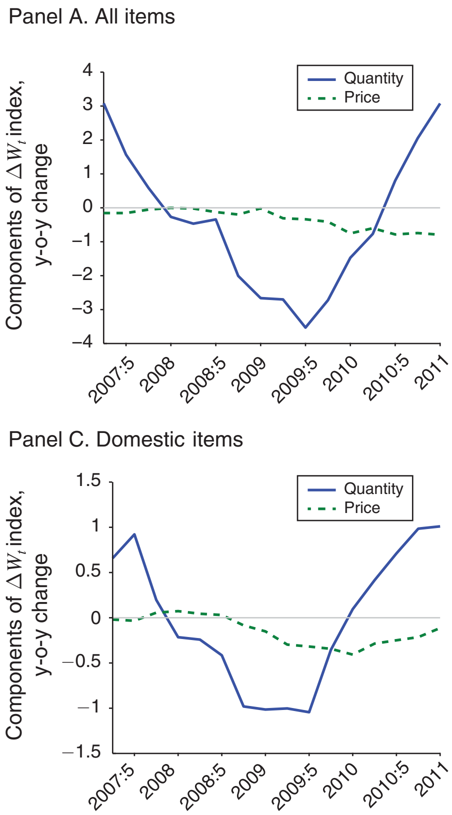

Figure 10. Changes in Prices and Consumed Quantities between High/Low Unit Value Items: All Items and Select Subsets

Notes: This figure reports price and quantity share indexes that measure y-o-y shifts in relative prices and consumed quantities between high/low unit value items within four-digit product groups. A negative value for the quantity index, defined as the quantity component of the index in (6), implies a systematic shift in quantities consumed toward lower unit value items. A positive value for the price index, defined as the price component of the index in (6), implies a systematic decease in the relative price of lower unit value items. The reported indexes aggregate results for product groups using each group's share in total F&B expenditures.

that a similar shift in consumed quantities toward less expensive items was present within imported items, within domestic items as well as within store products, for which we are not able to identify origin.[30] At the same time, we find no systematic shifts in the relative price of low/high unit value items (dashed line) over the boom-bust-recovery cycle. In particular, there is no evidence that during the crisis the shift in quantities toward lower unit value items was facilitated by a decrease in the relative price of lower unit value items, as would be captured by positive values of the price component in (6).

Shifts in Item Mix and Expenditure Switching.—The key question for this paper is whether the shift in the consumed item mix toward lower unit values during the crisis induced expenditure switching—a reallocation of expenditures from expensive

[30] We also find the same shift in item mix present when the data are split by the store type.

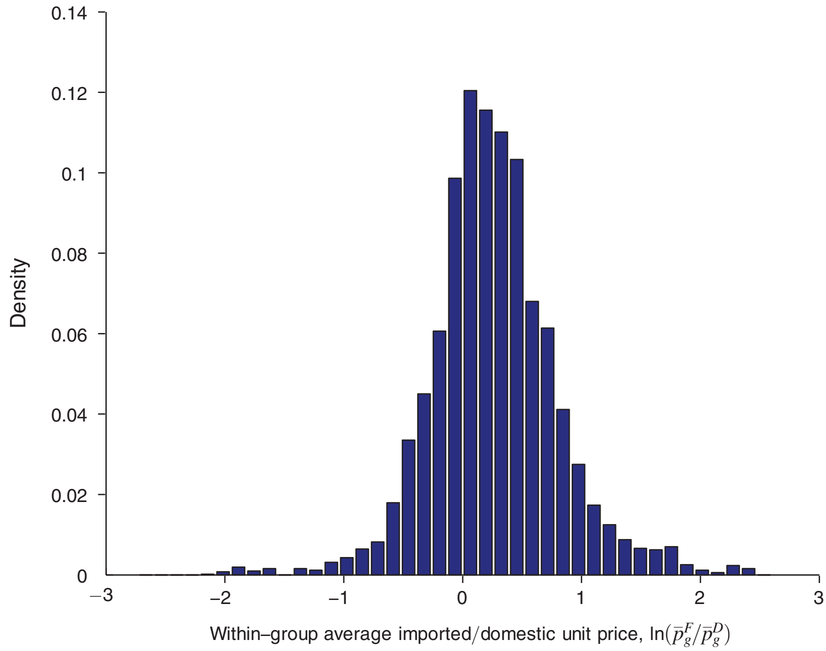

Figure 11. Distribution of Within-Product Group Average Unit Value of Imported Items Relative to Domestic Items at the Four-Digit Product Code Level

Notes: This figure plots the distribution of the average within-group unit values of foreign items relative to domestic items at the four-digit product code level. To construct the average relative unit value, we compute for each item in a product group the sample median relative unit value within a product group, $\overline{p}_{ig}$, based on available observations for $p_{igt}/\overline{p}_{gt}$, where p_{igt} is the unit value for item i in group g at time t and $\overline{p}_{gt}$ is defined below equation (6), and then compare the average unit value of imported and domestic goods in each product group, $\overline{p}_g^F/\overline{p}_g^D$.

foreign items to less expensive domestic ones. This section argues that this was indeed the case. We show that imported goods tend to be more expensive, which in combination with the documented shift toward lower unit values during the crisis induced the expenditure switching.

To compare price levels of domestic and imported goods, we compute for each item the sample median relative unit value within its four-digit product group, $\overline{p}_{ig}$, based on available observations for $p_{igt}/\overline{p}_{gt}$, and then compare the average unit value of imported and domestic goods in each product group, $\overline{p}_g^F/\overline{p}_g^D$. Figure 11 plots the distribution of the resulting foreign/domestic unit value differences across product groups. Imported items are on average 28 percent more expensive than their domestic counterparts in the median product group. The mean difference is 32 percent. These differences are persistent over time, with the mean difference varying between 26 percent and 37 percent.

Figure 11 also reveals sizable heterogeneity in the relative unit value of foreign to domestic items across product groups. Although imported goods are more expensive on average, the reverse is true for a fraction of product groups accounting for 19 percent of total expenditures in the dataset. This heterogeneity in average foreign/domestic unit values across product groups can be explored to link shifts in the consumed item mix with expenditure switching. In particular, we begin by computing the quantity component of the index in (6) separately for two data subsets: one

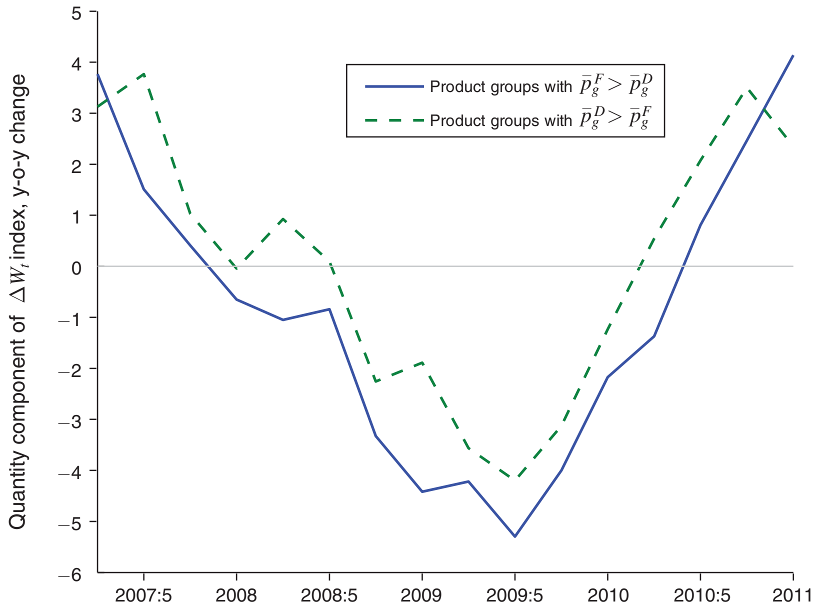

Figure 12. Changes in Consumed Quantities between High/Low Unit Value Items: Groups with Higher/Lower Average Domestic Unit Values

Notes: This figure reports a quantity component of the index in (6) separately for product groups with higher/lower average domestic unit values. A negative index value implies a systematic shift toward lower unit value items. The reported indexes are an expenditure weighted average of product group indexes.

containing product groups where imported items are on average more expensive, and the other containing product groups where domestic items are on average more expensive. Figure 12 shows that quantities consumed shifted toward lower cost unit value items in both data subsets.[31] For the subset of product groups where imports are more expensive (the solid line), which represents 81 percent of total expenditures, this finding suggests that consumers re-allocated expenditures from more expensive imports to less expensive domestic goods.

However, the item-level index in (6) leaves open the possibility that the switching within product groups is taking place within domestic and within imported items rather than across the two item types. To address this concern, we impose within-group source homogeneity and aggregate domestic and imported items in each product group into two composite goods—one containing all domestic items and the other containing all imported items—and then recompute the index. Formally, the quantities and prices of the two composite goods are computed as $q_{gt}^j = \sum_{i \in I_{gt}^j} q_{igt}$ and $p_{gt}^j = \sum_{i \in I_{gt}^j} p_{igt} q_{igt} / \sum_{i \in I_{gt}^j} q_{igt}$, where $j = \{D, F\}$ denotes the subsets of domestic and imported items in product group g. An index analogous to (6) is then constructed as

$$\Delta W_{gt}^{F/D} = \sum_j \left(\ln \varphi_{gt+1}^j - \ln \varphi_{gt}^j\right) \left(\ln p_{gt}^j - \ln \bar{p}_{gt}\right), \tag{7}$$

[31] The price change components (not shown) are close to zero.

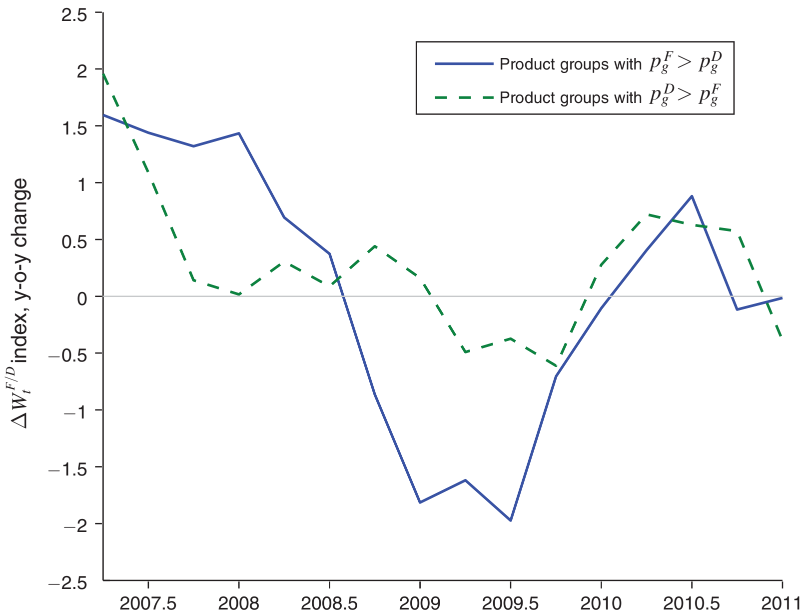

Figure 13. Expenditure Switching between Composite Foreign and Domestic Goods

Notes: This figure reports an expenditure share index that measures y-o-y expenditure switching between composite foreign and domestic goods within four-digit product groups. Index values are reported separately for (i) groups where the unit value of the composite foreign good exceed the unit value of the domestic counterpart, and (ii) groups where the unit value of the composite domestic good exceed the unit value of the foreign counterpart. Group indexes are aggregated using expenditure shares. A negative index value for both indexes implies the shift of expenditures towards the less expensive composite good.

where $\varphi_{gt}^j \equiv p_{gt}^j q_{gt}^j / \sum_{i \in I_{gt}^j} p_{gt}^j q_{gt}^j$.[32]

The index in (7) eliminates all sources of switching between items, except for switching that takes place between the two composite (domestic and imported) goods. Consequently, when applied to product groups where the composite imported good is relatively more expensive than its domestic counterpart, a negative index value implies expenditure switching toward the composite domestic good.[33]

Figure 13 plots two indexes based on (7), where the solid line denotes the index for product groups with a more expensive foreign composite good, and the dashed line denotes the index for product groups with a more expensive domestic composite good. As can be seen, these composite indexes follow the same pattern as the item-level indexes in Figure 10 and Figure 12, exhibiting negative values during the crisis. Thus, even when the switching is limited to shifts across imported/domestic goods, consumers substituted toward less expensive goods during the crisis. For product groups where foreign goods are more expensive, Figure 13 implies

[32] We do not show the quantity and price components of this index separately, because Section IIB already documents the (lack of) relative price adjustment for imported goods. Results for a quantity share index are almost identical to the findings for the expenditure share index.

[33] To construct this index series, we restrict the dataset to four-digit product groups that have at least one domestic and foreign item in each period, and eliminate extreme growth rates in the constructed expenditure shares for the two composite goods, which can generate noise in the series given the level of aggregation.

expenditure switching. For the product groups where domestic goods are more expensive, the shift in expenditures toward lower unit value imported goods is somewhat less pronounced. This more subdued *reverse* expenditure switching can in part be explained by the reduced scope for savings from switching expenditures in product groups where domestic goods are on average more expensive. In such product groups, the domestic composite good is 20.6 percent more expensive (in the median group), while for product groups where imported goods are more expensive, the median price differential is 49.6 percent.

To map the findings based on the two types of product groups in Figure 13 into the traditional measure of expenditure switching, as depicted in Figure 5, we compute total expenditure switching for each type of group ($\overline{p}_g^F > \overline{p}_g^D$ and $\overline{p}_g^D > \overline{p}_g^F$), and find that the switching from higher unit value imported items to lower unit value domestic items dominates. During the crisis period, aggregate expenditure switching is generated entirely by groups in which imported goods are on average more expensive, accounting for 105 percent of aggregate expenditure switching in Figure 5. Meanwhile, the import share actually increased in product groups where domestic goods are more expensive, amounting to a -5 percent contribution to the aggregate expenditure switching from foreign to domestic goods.

Our finding that there is switching toward items with lower unit values during crises is consistent with earlier findings by Burstein, Eichenbaum, and Rebelo (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-level data in the United States has noted that consumers search for cheaper goods by switching stores during recessions (Coibion, Gorodnichenko, and Hong 2015), as well as differences in consumption across cities and household income levels (Handbury 2013), but no one has examined the international dimension yet.

In sum, the empirical findings presented in this section provide evidence that expenditure switching in Latvia was primarily driven by substitution between domestic and foreign goods within detailed product groups. However, this switching was not accompanied by a corresponding relative price change. Rather, we provide evidence that consumers switched between cheap and expensive goods as income fluctuated, and in particular that they substituted from expensive foreign goods to cheaper domestic ones during the crisis.[34] One remaining issue with the findings reported in this section is that the constructed expenditure and quantity share indexes do not allow for the interpretation of the economic magnitudes of the expenditure switching that our proposed channel can account for. This issue is addressed in the following section.

[34] We have also examined an alternative approach at capturing shifts in item mix within product groups, based on a decomposition of the change in product group's average unit value into contributions from underlying changes in item prices and quantities (Boorstein and Feenstra 1987). Findings from this methodology are similar to results in Section IIC.

III. Demand Model Estimation

To formally quantify the importance of relative prices and income on expenditure switching, we model a consumer's expenditure allocation for F&B to provide structure for an estimation strategy, which exploits the dataset at the item level. We follow the literature that uses scanner-level 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.[35]

Given the documented heterogeneity in unit values 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's (2006) 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 nonhomotheticity at the bottom layer of the utility function.[36] 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 strategy captures an important potential channel that we wish to test; i.e., that consumers substituted to cheaper goods (within a product group) during the crisis, irrespective of relative price changes.

Online Appendix C provides the setup and solution to the consumer's allocation problem, and derivation of the within-group expenditure share, which is a function of relative prices and a quality parameter. The solution does not specify a function for how income affects a consumer's allocation of expenditures, but to provide an estimation strategy, we follow Hallak (2006) and assume that the consumer's intensity for demand of an item's quality in a given group is a linear function of aggregate expenditures. Given this assumption and the model solution, we can write the log-difference of item i's within-group g expenditure share at time t, φ_{igt}, as

$$\Delta \ln \varphi_{igt} = \Delta \ln N_{gt} + (1 - \sigma_g) \Delta \ln\left(\frac{p_{igt}}{P_{gt}}\right) + (\sigma_g - 1)\,\mu_g \ln \theta_{ig} \Delta \ln C_t, \tag{8}$$

where N_{gt} is the total number of items in product group g; σ_g is the elasticity of substitution between items in group g; p_{igt} is an item i's price level; P_{gt} is group g's price index; μ_g is a group g specific coefficient that captures the importance of aggregate expenditures on the intensity of demand of an item's quality, denoted by θ_{ig}; and C_t is aggregate expenditures.

[35] See Broda and Weinstein (2010) or Handbury (2013) for recent contributions using nested utilities and scanner-level data. Blackorby, Boyce, and Russell (1978) is an early contribution that uses nested utility, and which also allows for nonhomothetic preferences.

[36] Hallak (2006) takes the supply of quality and income as exogenous in a partial equilibrium setting, like ours. See Feenstra and Romalis (2014) for a general equilibrium model, where quality is an endogenous outcome. Furthermore, Choi, Hummels, and Xiang (2009) and Fajgelbaum, Grossman, and Helpman (2011) study how countries' income distributions affects trade and quality in a more general setting.

A. *Estimation Strategy*

We operationalize the estimation of (8) using the following baseline linear regression:

$$(9) \qquad \Delta \ln \varphi_{igt} = \alpha_{gt} + \beta_{1g} \Delta \ln\left(\frac{p_{igt}}{P_{gt}}\right) + \beta_{2g} \ln \overline{p}_{ig} \Delta \ln C_t + \varepsilon_{igt},$$

where α_{gt} is a four-digit product group $\times$ time fixed effect, which absorbs all explanatory variables that only vary in the *gt* dimension; $\beta_{1g} \equiv 1 - \sigma_g$; $\beta_{2g} \equiv \mu_g(\sigma_g - 1)$; $\overline{p}_{ig}$ is a proxy for the quality parameter θ_{ig}, and is calculated as the sample median of each item's relative unit value within a product group, $p_{igt}/\overline{p}_{gt}$, where the group's average unit value $\overline{p}_{gt}$ is defined below (6), and ε_{igt} is a random disturbance term. As in the model, we interact the *quality* term $\ln \overline{p}_{ig}$ with the growth rate of income, $\Delta \ln C_t$.[37] We use quarterly real per-capita household consumption for C_t, which we take from the International Financial Statistics (IMF).

We estimate (9) using the same data sample as described in Section II, though we drop four four-digit product groups that do not contain enough data to identify the coefficients of interest. The final regression sample comprises of 372,484 store item $\times$ time observations, and 387 product groups.[38] We estimate two versions of the baseline regression (9). The first model, which we call the CES model, restricts all β_{2g} to zero. Therefore, only changes in relative prices will affect an item's share. The second model, which we call the NH (nonhomothetic) model, runs (9) unrestricted.

The estimating equation (9) is based on a partial equilibrium demand model, which treats several variables as given, and ignores potential shocks that may impact both the quantity and the price of goods, thus inducing a correlation between ε_{igt} and price changes. Furthermore, the aggregate price indexes are model-based, and therefore a function of some of the parameters we wish to estimate. To address these issues, our estimation strategy follows a multi-pronged approach, which addresses: (i) biases due to measurement error induced by deviations from model-based price indexes; (ii) supply shocks that may have impacted the relative supply of domestic and imported goods; (iii) price discrimination by firms at the item level; and (iv) omitted variable bias. Our strategy relies on different fixed effect configurations, as well as two instrumental variables estimation approaches. All details are described in online Appendix D.

B. *Estimation Results*

Table 6 presents baseline results for the estimated heterogeneous coefficient regression model (9). The estimators presented are weighted means of the estimated coefficients, where the weights are based on a product group's average expenditure

[37] We use the terminology of *income* rather than *expenditure* in what follows for ease of exposition.

[38] 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. Furthermore, we also run regressions dropping the alcoholic beverage product groups. Results were robust to these restrictions, and are available upon request.

Table 6—CES and Nonhomothetic Models' Heterogeneous Coefficient Regressions: Weighted-Mean Estimates

	(1)	(2)
$\Delta \ln(p_{igt}/P_{gt})$	−2.938	−2.949
	(0.028)	(0.027)
$\ln \overline{p}_{ig} \times \Delta \ln C_t$		1.701
		(0.140)
Observations	372,484	372,484
Group × time FE	7,344	7,344
R^2	0.143	0.146

Notes: This table presents weighted means of the coefficients of regression model (9). The weights are a product group's share of total expenditures over the sample period. Column 1 presents the price coefficients for the CES model, and column 2 presents the price and income coefficients for the nonhomothetic model. These specifications are run with product group × time fixed effects. Standard errors clustered at the item level are in parentheses.

share over the sample.[39] Column 1 presents results for the CES model, where we only consider relative price changes and ignore potential income effects. The estimated relative price coefficient is -2.938, and is significant at the 1 percent level. This coefficient implies a price elasticity, σ, equal to 3.938.[40] Column 2 next presents the baseline results for the NH model. The estimated β_1 coefficient remains virtually unchanged and the estimated β_2 coefficient is positive and significant, with a value of 1.701, which implies a value of μ equal to 0.579.

Online Appendix D expands the estimation strategy along several lines, including more restrictive sets of fixed effects, nonlinear effects, instrumental variables, and numerous sample splits and other robustness checks (see online Appendix Tables A1–A4). The core results of this section are robust to all these modifications.

C. *Predicted Aggregate Expenditure Switching*

This section explores the quantitative importance of the estimated price and income effects in driving aggregate expenditure switching. Using the item-level data and the baseline estimated heterogeneous parameters for the CES and NH models, as summarized in Table 6, we predict each item's share in a given product group, and aggregate all predicted shares of imported goods in order to calculate the predicted import share in total F&B expenditures, and the corresponding expenditure switching for all sample periods, which we compare to the data. We focus on within-group expenditure switching. 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

[39] We choose to present weighted-means rather than simple means, since the predicted aggregate expenditure switching in Section IIA is also based on product group weights and heterogeneous coefficients.

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

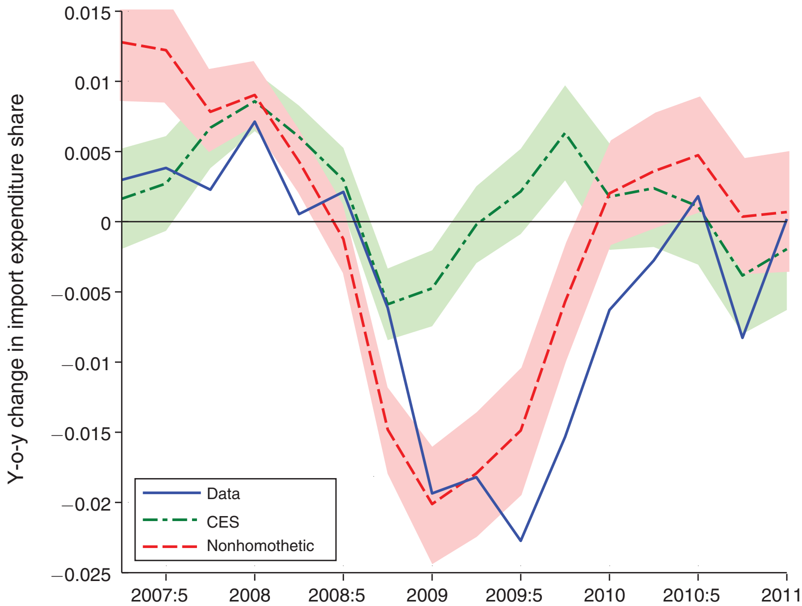

Figure 14. Model Estimated and Actual Within-Components of Expenditure Switching

Notes: This figure plots the within-component of expenditure switching observed in the data and estimated using the model based on (9), for the CES and nonhomothetic models. The shaded areas are two standard error bands, calculated analytically based on clustered standard errors at the grouptime level.

of the items' shares. Online Appendix E outlines all the details of the aggregation exercise.

Figure 14 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 Figure 5, 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 14, the predicted expenditure switching for the CES model (dashed-dotted 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, the nonhomothetic model's predicted values (dashed line) appear to track the data better throughout the sample, and match the switching during the precrisis boom period, as well as the switching during the crisis. In terms of quantities, the year-on-year within component of expenditure switching at its peak in 2009:III shows the import share falling by 0.022. For the same period, the CES model predicts an increase in the import share by 0.002, while the nonhomothetic model predicts a fall in the import share by 0.015. 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 nonhomothetic model is able to explain around two-thirds of what is observed in the data during the crisis.

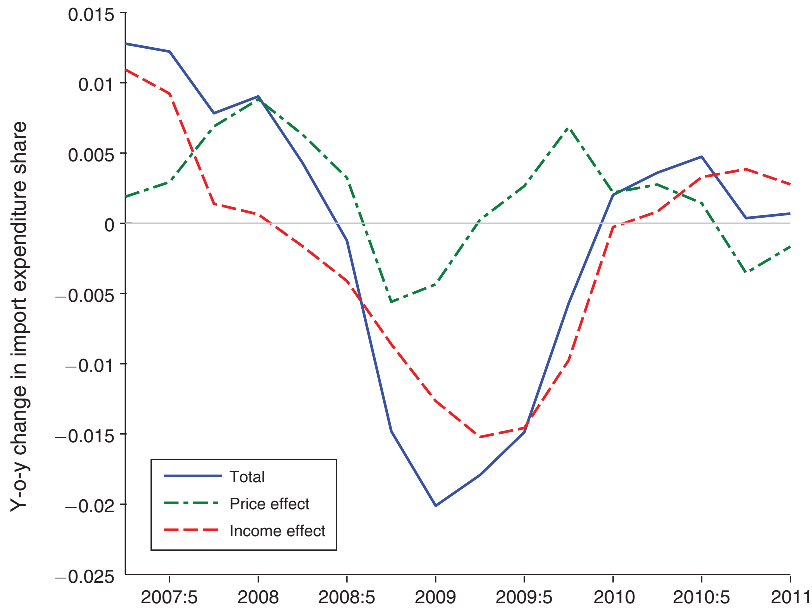

Figure 15. Nonhomothetic Model's Within-Components of Expenditure Switching: Income and Price Effects

Notes: This figure plots the estimated within-component of expenditure switching predicted by the nonhomothetic model, breaking it down into contributions due to (i) a price effect, and (ii) an income effect. The estimates are based on the full model (9).

Figure 15 decomposes predicted expenditure switching of the nonhomothetic model into separate price (dashed-dotted line) and income (dashed line) effects. We calculate these by either shutting down the price effect ($\beta_{1g}\text{s} = 0$), or the income effect ($\beta_{2g}\text{s} = 0$), 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.

IV. Conclusion

This paper measures what drove expenditure switching in Latvia during a sudden stop episode in 2008–2009, using a supermarket scanner-level dataset. 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 nonhomothetic 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 focuses on substitution between domestic and imports goods in a particular sector of the economy and for a country that maintained a peg during the crisis. Future work should investigate how relevant this nonhomothetic channel is in a more general setting, which incorporates exports and other sectors of the economy, as well as study how results vary across exchange rate regimes. Furthermore, this paper remains silent on several issues that are left for future work using the scanner-level dataset, possibly combined with micro supply-side data. There is need for a better understanding of what drove changes (or lack there of) in the relative price of domestic and foreign goods following Latvia's internal devaluation. Furthermore, what store-level maximization behavior would rationalize the results found in this paper given the macroeconomic situation faced by Latvia? Answering such questions, as well as gaining further insight into the price dynamics, is important to better understand the implications of policy prescriptions such as internal devaluations.

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