Unequal expenditure switching: Evidence from Switzerland

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Abstract

What are the unequal effects of changes in consumer prices on the cost of living? In the context of changes in import prices (driven by, e.g., changes in trade costs or exchange rates), most analyses focus on variation across households in initial expenditure shares on imported goods. However, the unequal welfare effects of non-marginal foreign price changes also depend on differences in how consumers substitute between imported and domestic goods, on which there is scant evidence. Using data from Switzerland surrounding the 2015 appreciation of the Swiss franc, we provide evidence that lower-income households have higher price elasticities. We quantify the contribution of heterogeneous elasticities for the unequal welfare effects of observed price changes between 2014–15 and for counterfactual shocks to the mean and dispersion of import price changes.

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1 Introduction

What are the unequal effects of changes in consumer prices on the cost of living? In the context of changes in prices of imported goods (due to, e.g., changes in trade costs or exchange rates), most attempts to answer this question have focused on variation across households in initial expenditure shares on imported goods; see, e.g., Fajgelbaum and Khandelwal (2016), Cravino and Levchenko (2017), and Borusyak and Jaravel (2021). However, the unequal welfare effects of non-marginal foreign price changes also depend on differences in how consumers substitute between imported and domestic goods (unequal expenditure switching), on which there is scant evidence. As noted by Deaton (1997, page 187):

Since my main interest here is in the distributional effects of price changes..., these [second-order] effects will change the conclusions only to the extent that the elasticities ... differ systematically between poor and rich. Although there is no reason to rule out such effects a priori, there is no reliable evidence on the topic.

In this paper, we document large differences in price elasticities across the income distribution and quantify their contribution to the unequal welfare effects of import price shocks.

In Section 2, we begin by describing the setting in which we measure both initial expenditure shares on imported goods across households as well as unequal expenditure switching. We focus on Switzerland in a period surrounding the abrupt appreciation of the Swiss franc on January 15, 2015, which markedly reduced import relative to domestic retail prices; see Figure 1. We measure initial import exposure across the income distribution using data on expenditure shares by household income groups across 296 consumption categories (using the Swiss Household Budget Survey) and import shares across 217 slightly more aggregated consumption categories (using the disaggregated data underlying the Swiss CPI). To study unequal expenditure switching, we turn to higher-frequency and more detailed barcode-level Swiss Nielsen Homescan data, cov-

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Footnotes:

1 See also Friedman and Levinsohn (2002), Porto (2006), and Carroll and Hur (2020).

2 The Swiss National Bank (SNB) adopted a minimum exchange rate of 1.20 CHF per EUR in September 2011. Developments abroad in late 2014 and early 2015 prompted the SNB to unexpectedly abandon this policy on January 15, 2015. The subsequent appreciation episode came after a period of remarkable exchange rate stability, was significant (the EUR/CHF appreciated by 14.7 percent by the end of June), and—in contrast to many episodes with large swings in international relative prices—occurred against the backdrop of stable economic aggregates and nominal income inequality in Switzerland. Import prices fell by more at the border than at the retail level (Auer et al., 2021); we focus on the latter given our emphasis on expenditure switching and welfare at the consumer level.
In response to the 2015 CHF appreciation, the import share within the Homescan data rises and this increase is greater for lower-income households; this differential change in import shares across incomes is not driven by a larger decline in import prices for lower-income households. Finally, we document that the 2015 CHF appreciation was accompanied by an increase in the dispersion of price changes within the set of imported goods.

In Section 3, we characterize a set of sufficient statistics to answer our motivating question: What are the unequal welfare effects of changes in consumer prices (more specifically, changes in import relative to domestic prices in our quantitative applications) through their impact on the cost of living? A large literature has addressed this question by applying a first-order approximation of the expenditure function and, hence, focusing on variation across households in initial expenditure shares on different goods. We apply known results in microeconomic theory—see, e.g., Hausman (1981) and more recently Baqee and Burstein (2023)—to provide an exact answer to this question for non-marginal

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3 In terms of measuring initial import shares across the income distribution, our paper is most closely related to Borusyak and Jaravel (2021), who also use detailed data on consumer expenditures and import shares across the full economy rather than for aggregate industries and directly observe household-specific import shares on consumer packaged goods (and motor vehicles).
price changes, taking expenditure switching into account. The sufficient statistics to calculate a household’s compensating variation in response to a given change in income and prices are initial expenditure shares across products and compensated cross-price elasticities (i.e. cross-price elasticities along the initial indifference curve).\footnote{There are well known price indices (e.g. Törnqvist or Sato-Vartia) that incorporate information on observed expenditure shares over time and—under strong assumptions—measure welfare beyond first-order approximations. For recent alternative approaches, see Atkin et al. (2020), Jaravel and Lashkari (2022), and Baqaae et al. (2022). In contrast to all of these approaches, we calculate changes in welfare in response to counterfactual changes in prices and income; see Section 5.} The unequal effects of price changes on the cost of living are shaped by differences in initial expenditure shares and differences in compensated cross-price elasticities, and these effects increase in the dispersion of price changes across goods. In practice, estimating cross-price elasticities (compensated or uncompensated) between all goods in the economy is infeasible without additional assumptions. We therefore impose nested, generalized non-homothetic CES preferences, building on Matsuyama (2019), Fally (2022) and Comin et al. (2021). Income elasticities can be non-unitary; and elasticities of substitution between goods within a sector can vary between indifference curves, but (as in standard trade models) are constant along any indifference curve.

In Section 4 we estimate how elasticities of substitution between goods in the Homescan data vary with income, taking two approaches that leverage distinct sources of variation. In our first approach, we use variation in changes in import relative to domestic expenditures between 2014 and 2015 (surrounding the appreciation) across higher- and lower-income households. Our identification assumption is that import demand shocks in 2015 are not systematically different across incomes. In our second approach, we use variation in changes in expenditures across individual barcode products and variation in product price changes. In this case, we control for product-specific demand shocks and household-specific import demand shocks. We instrument for the interaction between initial household income and the product-specific price change using an interaction between household income and a product “cost shifter.” Our cost shifter exploits variation across border groups—an aggregation of products—in the invoicing currency of imports. Specifically, we measure the share of imported goods in each border group that are denominated in EUR, using information from the goods-level survey underlying the calculation of the official Swiss import price index. Because of the stickiness of import prices at the border in their invoicing currency, Swiss retail prices of imported goods are more responsive to the appreciation if imports are denominated in EUR rather than in CHF; see Auer et al. (2021). Given additional controls and our instrumentation strategy, the exclusion restriction in the second approach is substantially weaker than in the first approach.
In spite of these differences, we obtain very similar quantitative results across approaches. The elasticity of substitution between goods in the Homescan data is substantially lower for higher-income households: for example, the difference between the elasticities of substitution between two households, where one has an income three times that of the other, is 2.4 under the first approach and 2.1 under the second approach. These approaches identify differences in elasticities across incomes. To estimate the level of these elasticities, we make stronger identification assumptions, and the resulting estimates vary more across the two approaches. However, in our analytic and quantitative results we show that, conditional on differences in price elasticities across households, the unequal welfare effects of price changes are not very sensitive to elasticity levels.

Our estimates of higher price elasticities for lower-income households are qualitatively consistent with demand system estimates in industrial organization—see, e.g., Berry et al. (2004) among many others—and findings on shopping behavior in macroeconomics—see, e.g., Kaplan and Menzio (2016) and Aguiar and Hurst (2007)—all of which support Harrod’s (1936) Law of Diminishing Elasticity of Demand, which postulates that demand elasticities decrease in income. In the spatial economics literature, there is scant and conflicting evidence on differences in elasticities across household incomes. Argente and Lee (2021) and Faber and Fally (2022) find very small differences. Handbury (2021) finds differences in elasticities across incomes very similar to our estimates when controlling for the fact that higher-income households have a greater willingness to pay for quality but finds opposite results when not; our approach incorporates these differences in willingness to pay for quality (by incorporating household × barcode product-specific demand shifters that cancel out when estimating demand elasticities using changes rather than levels of household expenditure shares by income). Moreover, whereas these papers use either Hausman instruments or the approach developed by Feenstra (1994) and Broda and Weinstein (2006), we exploit exogenous variation in price responsiveness to an exchange rate shock.

Finally, in Section 5 we quantify how differences in price elasticities estimated using

\[5\] Bems and di Giovanni (2016) document that a large aggregate decrease in income in Latvia reduced import shares (since high-quality imports are more income elastic), and Coibion et al. (2015) show that households switch expenditures toward low-price stores when local economic activity falls. Rather than focusing on expenditure switching due to changes in income, we focus on heterogeneous expenditure switching across the income distribution in response to changes in relative prices.

\[6\] While Handbury (2021) calculates regional differences in price indices abstracting from her estimated differences in price elasticities, we show that differences in elasticities are quantitatively important in shaping the unequal welfare effects of large price changes.

\[7\] When we use a Hausman instrument, we find small differences. In Appendix B.2 we show that the Hausman instrument may be endogenous in our Swiss context (where there is little spatial variation in price changes).
the Homescan data shape the unequal welfare effects of changes in prices. For these exercises, we consider households at three distinct income levels ranging from 20,000 to 120,000 CHF per year with respective elasticities of 6.6, 4.4, and 3. We first calculate changes in the welfare-relevant price index over grocery products in the Homescan data in response to observed price changes between 2014 and 2015, following the CHF appreciation. Prices fell on average by 1.2% between 2014 and 2015. The welfare-relevant price index (taking expenditure switching into account) declined by 1.6% for households with income of 120,000 CHF and by 2.2% for households with income of 20,000 CHF. This gap between income groups is accounted for by unequal expenditure switching between imports and domestic goods as well as between products with different price changes within the set of imports and domestic goods. If price elasticities were equal across income groups, then the gap in price indices would be small and of the opposite sign.

We then calculate changes in welfare in response to counterfactual changes in import relative to domestic prices, considering, first, only uniform price changes within each set of goods and, second, incorporating variation in price changes within each set of goods. To conduct these counterfactuals, we use our measures of import shares for each income across all consumer goods and we impose that our estimated differences in price elasticities across incomes within our Homescan data apply more broadly. To highlight the nonlinearities from expenditure switching, we consider import price shocks that are larger than the one induced by the 2015 CHF appreciation. Uniform import price increases harm higher-income households more than lower-income households in Switzerland for two reasons: (i) higher-income households have higher initial import shares (since they spend relatively more on non-grocery goods, which are more tradable than groceries) and (ii) they substitute away from imported goods less. For large changes in prices, the impact of unequal expenditure-switching on welfare is substantial. For example, consider

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8Bai and Stumpner (2019) and Jaravel and Sager (2019) construct income-group and product-category-specific inflation rates and project these on changes in import penetration induced by China. Hottman and Monarch (2020) focus on differences in import price inflation rates across US households. Relative to these papers, we estimate differences in import elasticities across income groups, which we then use to quantify welfare changes for observed and counterfactual price shocks.

9Given our focus on the expenditure-side effects of foreign price shocks, in the counterfactuals we abstract from changes in the income distribution. There is a large empirical and theoretical literature on the impact of international trade on income inequality with multiple factors; see e.g. Burstein and Vogel (2017), Cravino and Sotelo (2019), Galle et al. (2022), and Adao et al. (2022). See, e.g., He (2018) and Borusyak and Jaravel (2021) for papers incorporating both income and expenditure-side inequality induced by trade.

10Variation of import shares with household income differs across countries depending, among other things, on whether the country is high or low income and has a comparative advantage in goods with high- or low-income elasticities. For example, Borusyak and Jaravel (2021) document that imports are flat throughout the income distribution in the US. Therefore, if we applied our counterfactuals to the US context, unequal expenditure switching would be the only channel inducing unequal welfare effects (via variation in cost of living).
a 20% uniform increase in import prices relative to domestic prices, which is not uncom-
mon in the context of large exchange rate devaluations; see, e.g., Burstein et al. (2005) and
Cravino and Levchenko (2017). In response to this shock, welfare of a household with an-
nual income of 20,000 CHF falls by about one third less than for a household with annual
income of 60,000 CHF. Almost half of this differential is accounted for by differences in
price elasticities. In practice, the variance of price changes within imports and domestic
goods rose during the 2015 CHF appreciation. If in our counterfactuals we additionally
consider an increase in price dispersion within imported and domestic goods, then the
importance of unequal expenditure switching can be substantially larger.

In summary, we make four main contributions. First, we document heterogeneity
across incomes in expenditure switching in response to an exogenous shock to import
prices. Second, we estimate differences in elasticities of substitution across incomes using
plausibly exogenous variation resulting from this shock to import prices and its hetero-
geneity across product groups. Third, we apply results in welfare economics to charac-
terize differential welfare changes due to heterogeneous price elasticities in response to
foreign-induced price changes. Fourth, we show quantitatively that unequal expenditure
switching contributes substantially to differences in welfare across the income distribu-
tion in response to observed Swiss price changes and to plausible counterfactual shocks
to the mean and dispersion of import price changes.

2 Data and stylized facts

2.1 Data

In this section we provide an overview of the main data sets employed in the paper.
Additional details and data sources are provided in Appendix A.

AC Nielsen Homescan data and import status. AC Nielsen Homescan, Nielsen Switzer-
land (2016), includes information on household (HH) characteristics and shopping trans-
actions of a demographically and regionally representative sample in Switzerland during
the period surrounding the 2015 appreciation: January 2013 to December 2016. The data
includes approximately 3,300 households in 2014.

Participating households record purchases—of food, beverages, personal care (health
and beauty aids), and other selected general merchandise—in supermarkets and drug-
stores (we refer to these goods as groceries). Individual products are identified by their
barcode (European Article Number or EAN, which we often refer to as a product). In the
appendix, we describe a number of adjustments we make to the data, such as dropping
likely recording errors by households. In the raw data, an observation is a transaction that includes the household identifier, EAN code, quantity purchased, price paid (net of good-specific discounts due to e.g. coupons), date of the shopping trip, and the name of the retailer. We drop all transactions that occur abroad. See Burstein et al. (2022) for an analysis of cross-border shopping in Switzerland using the Homescan data. We measure each product’s price (in logs) as an average of transaction-level log prices in the corresponding time period, weighing transactions by expenditures.

The Homescan data come with a rich set of socioeconomic characteristics for each household, summarized in Table A2 in Appendix A for the year 2014, including the two-digit zip code of residence, the educational level of the household’s main earner, the number of household members (and the number of those under 10 and over 70 years old), and total household pre-tax annual income reported in seven bins. Given each of these characteristics, we infer a level of household pre-tax income in 2014 for each Homescan household using additional data from the Swiss Household Panel (FORS). We do so by projecting the level of 2014 pre-tax household income in FORS on a set of household characteristics available both in FORS and Homescan, including indicator variables for the seven pre-tax income bins available in Homescan (and which can be constructed in FORS). We then predict household income in Homescan using these coefficients.

We augment the Homescan data with information on whether individual products are imported or produced domestically. Whereas EANs provide information on the country in which a product has been registered, in many instances this is not the country in which the product has been produced. However, that information is disclosed on the label of each product. We use the label information that Auer et al. (2021) collected from codecheck.info. Coverage is not complete and notably excludes most fruits and vegetables, EANs assigned by store managers locally, and goods that are only occasionally sold in grocery stores such as toys, clothing, or household electronics. Our measure of import status for each individual product is fixed over time, as obtained from codecheck.info between October 2015 and March 2016. We drop products for which import status is unknown.

Comparing columns 1 and 2 of Table A1 in Appendix A, we see that out of 69,088 unique goods and expenditure of CHF 11.1 million, there are 8,409 unique goods and expenditure of CHF 4.2 million with known import status; the share of expenditure for

\[\text{Roughly 90\% of imported goods in our data are from the European Union. Our results are robust to dropping imports from other origins.}\]
which the production location is known is approximately 38%.\textsuperscript{12,13} We further divide products with known import status into imports and domestically produced goods in columns 3 and 4. A similar number of unique imported and domestically produced goods were purchased and the import share (at retail prices) of expenditures was 26.9%.

**Expenditure and import shares by income group and sector.** To calculate expenditure shares and import shares by income group across each consumption category, we use two data sets provided by the Swiss Federal Statistical Office (SFSO).

The first data set, the Swiss Household Budget Survey (HBS), reports information about consumption expenditures by income group and consumption category for the periods 2012–14 and 2015–18 based on roughly 250 households per month randomly selected from a large and representative registry. At the lowest level of disaggregation, there are 296 consumption categories for goods and services, such as “rice”, “pasta”, or “tickets for public transport.” The SFSO collects expenditures on these consumption categories separately for each of five income groups. We use data for the pre-appreciation period 2012–14 to construct sectoral expenditure shares for each group. While we construct these sectoral expenditure shares by income group for each of the 296 disaggregated consumption categories, Table 1 displays expenditure shares in the aggregate and separately for each income group, aggregated up to three sectors: groceries (matching as close as possible our Homescan goods), non-grocery goods, and services.

The second data set contains a cross section of import shares by disaggregated product category, obtained by the SFSO via firm surveys published in 2016 based on information from previous years. These shares, used by the SFSO to calculate a CPI for imported goods, are available at a similar disaggregation to the ones in the HBS data. We combine these import shares—which vary across disaggregated category—with the HBS data—which vary across disaggregated category and income group—to construct import shares by income group within each of our three aggregate sectors. To do so, we assume that different income groups have common import shares within each disaggregated product category, an assumption we do not impose in the Homescan data since we observe the import status of individual products. Table 1 displays the resulting import shares by

\textsuperscript{12}Many of the products for which we do not observe import status appear in the Homescan data for only a short period of time. If we keep only those products that were purchased at least once per year between 2013 and 2016, the share of expenditure on goods with known origin is 47% instead of the 38% we observe in our baseline sample.

\textsuperscript{13}One concern might be that expenditure on products for which we do not observe import status is correlated with household income. However, in Appendix A we show that household income is not significantly correlated with the household’s share of expenditure in 2014 on products for which we do not observe import status.
Table 1: Expenditure and import shares by income group and sector

<table>
<thead>
<tr>
<th>Annual income</th>
<th>Expenditure shares</th>
<th>Import shares</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Grocery</td>
<td>Other goods</td>
</tr>
<tr>
<td>– 60,252</td>
<td>20.1</td>
<td>18.5</td>
</tr>
<tr>
<td>60,252 – 88,032</td>
<td>18.6</td>
<td>21.6</td>
</tr>
<tr>
<td>88,032 – 119,736</td>
<td>18.0</td>
<td>23.4</td>
</tr>
<tr>
<td>119,736 – 164,244</td>
<td>17.1</td>
<td>24.3</td>
</tr>
<tr>
<td>164,244 –</td>
<td>15.1</td>
<td>25.6</td>
</tr>
<tr>
<td>All</td>
<td>17.2</td>
<td>23.5</td>
</tr>
</tbody>
</table>

Notes: Expenditure shares by income range and sector—aggregated to groceries, other non-grocery goods, and services—are from SFSO using the years 2012–14. Import shares are constructed from import shares in disaggregated product categories and expenditure shares by income across these categories. The final row represents the value of each column across all households and the final column represents the import share of each income group across all sectors.

Currency of invoicing of import prices at the border. Our instrument in Section 4 exploits variation across imported goods in the invoicing currency of prices at the border. We match individual barcode products in the Homescan data to groups of imported products at the border (border groups) and measure the share of imported products in each border group in 2014 that are denominated in EUR (out of those denominated in either EUR or CHF), using information from the goods-level survey underlying the calculation of the official Swiss import price index. For additional information, see Auer et al. (2021).

2.2 Stylized facts

In this section we present our five stylized facts (SFs), indicating in brackets for each fact the data set that we use to calculate it. Facts 2–5 use Homescan rather than SFSO data because the SFSO data are not available at annual frequency. In Appendix B.1 we provide additional information.

SF 1 (SFSO): Import shares before the 2015 CHF appreciation were higher among higher-income households.

The right-most column of Table 1 displays the aggregate import share—across all consumption categories—for each of the five income groups in the SFSO data (for the 2012–14 period). Higher-income households had higher aggregate import shares in Switzerland, with the share rising monotonically from 21% to 28% between the bottom and top income  

14 The aggregate import share for groceries in the SFSO sample (37.9%) is higher than in the Homescan data (26.9%). In Appendix A, we show that this is due to expenditures in the SFSO data being comparatively concentrated in categories with high import shares. Applying common expenditure weights across categories in the SFSO and Homescan data yields more similar aggregate import shares.
Figure 2: Aggregate import responsiveness and heterogeneity across incomes in Homescan data

Notes: Import shares by year aggregated across all households (All), households with incomes below our sample median (Low income), and households with income above our sample median (High income). Point estimates and ten percent confidence intervals for each income group and year are calculated by regressing separately across household \( h \) within the low-income group and the high-income group \( 100 \times \frac{X_{hM}}{X_{hM} + X_{hD}} \) on a constant, weighing \( h \) by \( X_{hM} + X_{hD} \) and clustering by income quantile (of which there are fifty).

groups in the SFSO data. This pattern is accounted for by two relationships. First, the import share was much higher for non-grocery goods (across all income groups) than for groceries or services, and higher-income households spent a higher share of their income in this sector. Second, higher-income households had a higher import share within the non-grocery goods aggregate sector and, to a lesser extent, within services.\(^{15}\)

On the other hand, import shares within groceries were not strongly correlated with income. This is evident in the SFSO data from Table 1. The same (weak) relationship holds within our product-level Homescan data, as shown in Table A5 in Appendix B.1.

SF 2 (Homescan): The import share increased following the 2015 CHF appreciation.

The aggregate import share in the Homescan data increased from 26.9% to 27.5% between 2014 and 2015, as shown in Figure 2. To show that this occurred within individual households—rather than from a change in the composition of expenditures across households—we regress each household’s import share of expenditure in each year on household effects and year effects, excluding the year 2014. These year effects identify the change, within households, in the import share of expenditures between each year and 2014. We find no economically or statistically significant differences between 2013 and 2014. On the other hand, within households the import share was higher in 2015 than it was in 2014 and continues increasing into 2016; see Figure A1 in Appendix B.1, which reports the rise in year effects estimated separately over all monthly time horizons within the year (we define horizon \( j \) as the first \( j \) months in year \( t \) and in 2014).

SF 3 (Homescan): Import shares increased less for higher-income households following the 2015 CHF appreciation.

\(^{15}\)This is largely because high-income groups have higher budget shares on cars and cars in Switzerland tend to be imported.
Table 2: Heterogeneous expenditure switching towards imports

<table>
<thead>
<tr>
<th></th>
<th>log(Income)</th>
<th>High income</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>Income 2013</td>
<td>-0.472*</td>
<td>-0.489**</td>
</tr>
<tr>
<td></td>
<td>[0.266]</td>
<td>[0.265]</td>
</tr>
<tr>
<td>Income 2015</td>
<td>-0.727**</td>
<td>-0.813***</td>
</tr>
<tr>
<td></td>
<td>[0.272]</td>
<td>[0.279]</td>
</tr>
<tr>
<td>Income 2016</td>
<td>-0.953***</td>
<td>-0.970***</td>
</tr>
<tr>
<td></td>
<td>[0.321]</td>
<td>[0.339]</td>
</tr>
<tr>
<td>Observations</td>
<td>11630</td>
<td>11630</td>
</tr>
<tr>
<td>Control size</td>
<td>X</td>
<td>X</td>
</tr>
<tr>
<td>All controls</td>
<td>X</td>
<td>X</td>
</tr>
</tbody>
</table>

Notes: Results of estimating $x_{hMt} / (x_{hMt} + x_{hDt}) = F_{Et} + F_{Eh} + \sum_{y=2014}^{t} (\beta_y \cdot inc_h + \zeta_t \cdot k_h) + \epsilon_{ht}$, where $x_{hMt}$ and $x_{hDt}$ are household $h$’s expenditure on imports and domestic goods in year $t$; $F_{Et}$ and $F_{Eh}$ are year and household fixed effects; $inc_h$ is either the logarithm of household income (columns 1–3) or an indicator that equals one if HH income is greater than the median in our sample of 57,647 CHF (columns 4–6); Income $y$ displays the coefficient on $\beta_y$. In columns 2 and 5, we control for HH size interacted with year. In columns 3 and 6 we additionally include an indicator for whether there is a child under 10 and an indicator if everyone in the HH is older than 70, each interacted with year. Robust standard errors are clustered by income quantile (of which there are fifty) and observations are weighted by the product of the number of households in the income quantile and the household’s share of expenditure in 2014 within its income quantile. *p<.1; **p<.05; ***p<.01

Figure 2 displays the import share by year separately for high- and low-income households (those with annual incomes above and below our sample median of approximately 57,600 CHF). Between 2014 and 2015 the import share of low-income households increased more than that of high-income households. The difference between the import shares of low- and high-income households expanded from approximately 0.5 to 1.1 percentage points between 2014 and 2015. This gap rose further to 1.2 percentage points in 2016. Figure 2 shows no evidence of pre-trends in the difference between low- and high-income import shares.

This greater increase in import shares among low-income households occurred within individual households, rather than from a change in the composition of expenditures across households. To show this, we estimate the differential change across household incomes in the import share of expenditures between each year and 2014 at the household level, controlling for household fixed effects. To measure income we use either the log of household income or an indicator for household incomes above the median.

Table 2 shows that import shares increased substantially less for high- than low-income households after the exchange rate appreciation. For example, a household with an income three times that of another experienced a roughly 0.7 percentage point smaller

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16 The higher import share for low-income households in 2014 displayed in Figure 2 appears inconsistent with Stylized Fact 1. However, the difference is not statistically significant, as revealed by the wide confidence in the figure and the results in Table A5.
increase in its import share between 2014 and 2015. This is not a continuation of pre-existing trends: the coefficients in Table 2 indicate that the gap between the import shares of low- and high-income households did not rise between 2013 and 2014 (see the first row of columns 1 and 4) before rising between 2014 and 2015. Changes between 2014 and 2015 also do not appear to be reversion to the mean: the coefficient in row 1 of column 4 is insignificantly different from zero and the coefficient in row 1 of column 1, while marginally significant, is much smaller than in row 2. These results are robust to including progressively more household-level controls interacted with year; see columns 2, 3, 5, and 6.

SF 4 (Homescan): The price of imported relative to domestic goods fell following the 2015 CHF appreciation. Neither import nor domestic price changes varied systematically with household income.

Regressing individual product prices on product fixed effects and month fixed effects (omitting a fixed effect for the month preceding the CHF appreciation) separately for imports and domestic goods, we find that import prices fell by approximately 2% relative to domestic prices following the appreciation (averaging the change between December 2014 and each month in 2015); see Figure 1 in the Introduction.

In Appendix B.1, we run a related regression that separates prices paid for each product by household income. We find no differential changes in import or domestic prices across income, either economically or statistically; see Figure A2 in Appendix B.1.\(^\text{17}\) This pattern is robust to further disaggregating prices by region in Switzerland and including region fixed effects in the regression. This implies that the pattern of heterogeneous changes in import shares across households with different incomes described in Stylized Fact 3 is not driven by lower-income households facing greater declines in relative import prices.

SF 5 (Homescan): The dispersion of price changes across goods was greater between 2014–15 than between 2013–14 and than between 2015–16, especially within the set of imported goods.

Table 3 displays the standard deviation across barcode products of annual log price changes within the set of imported goods, domestic goods, and all goods between years \(t\) and \(t + 1\) for \(t = 2013, 2014,\) and \(2015\). The standard deviation of log price changes rose from 5.8% to 7.6% for imported and from 4.2% to 4.8% for domestic goods between 2013–

\(^{17}\)We also find no relationship between income and the level of the price paid within individual barcode products. Relatedly, defining products as aggregations of barcodes, Handbury (2021) finds that most of the variation in prices paid across incomes is accounted for by differences in products purchased.
Table 3: Standard deviation of log price annual changes across barcode products

<table>
<thead>
<tr>
<th></th>
<th>Imports</th>
<th>Domestic</th>
<th>All</th>
</tr>
</thead>
<tbody>
<tr>
<td>2013–14</td>
<td>0.058</td>
<td>0.042</td>
<td>0.047</td>
</tr>
<tr>
<td>2014–15</td>
<td>0.076</td>
<td>0.048</td>
<td>0.058</td>
</tr>
<tr>
<td>2015–16</td>
<td>0.061</td>
<td>0.041</td>
<td>0.047</td>
</tr>
</tbody>
</table>

Notes: Expenditure-weighted standard deviation of annual log price change across barcode products between $t$ and $t+1$ for $t = 2013$, 2014, and 2015, for imported goods (column 1), domestic goods (column 2), and all goods (column 3).

14 and 2014–15; each approximately reverts to its initial level between 2015–16. We show that these patterns are robust to various alternative choices in Table A7 in Appendix B.1.

We will show that this increase in the dispersion of price changes—together with greater expenditure switching of lower-income households—contributes to the unequal welfare effects of the 2015 CHF appreciation.

3 Welfare impacts of price changes

Our objective is to construct a measure of the change in welfare for different households, starting from an initial observed period, in response to factual or counterfactual changes in income and in the prices of consumption goods. In Section 3.1 we provide sufficient statistics for this calculation under general preferences, building on results in micro theory. In Section 3.2 we restrict these preferences to a particular parametric form, which we use in the following sections to estimate differences in elasticities of substitution across households and to perform our quantitative applications.

3.1 General formulation

Household $h$’s preferences over $N$ consumption goods, indexed by $\zeta$ (taste shifters), can be represented by the expenditure function $e_h(p, u; \zeta)$, which indicates the minimum cost of achieving utility $u$ given a vector of prices $p$. The associated budget share on good $i$ is denoted by $b_{hi}(p, u; \zeta)$, which by Shephard’s lemma equals $\partial \log e_h(p, u; \zeta) / \partial \log p_i$. Given income $I$ (which we assume is equal to expenditures), the indirect utility function is $v_h(p, I; \zeta)$.

We consider a change in household $h$’s income from $I_{ht0}$ to $I_{ht1}$, prices, from $p_{ht0}$ to $p_{ht1}$, and taste shifters from $\zeta_{ht0}$ to $\zeta_{ht1}$. Our welfare measure is the compensated variation
evaluated at initial preferences, \( CV_h \), which is implicitly defined by

\[
v_h(p_{ht_0}, I_{ht_0}, \zeta_{ht_0}) = v_h(p_{ht_1}, e^{-CV_h I_{ht_1}}, \zeta_{ht_0}).
\]

In words, \( CV_h \) is the reduction in income (in logs) under the final budget set that makes the household with initial preferences equally well-off as under the initial budget set. Household \( h \) is better off under the final than initial budget set if and only if \( CV_h > 0 \).

We can express \( CV_h \) using the expenditure function as

\[
CV_h = \log \left( \frac{e_{ht}(p_{ht_1}, v_h(p_{ht_1}, I_{ht_1}, \zeta_{ht_0}); \zeta_{ht_0})}{e_{ht}(p_{ht_0}, v_h(p_{ht_0}, I_{ht_0}, \zeta_{ht_0}); \zeta_{ht_0})} \right) = \log \left( \frac{I_{ht_1}}{I_{ht_0}} \right) - \log \left( \frac{e_{ht}(p_{ht_1}, u_{ht_0}, \zeta_{ht_0})}{e_{ht}(p_{ht_0}, u_{ht_0}, \zeta_{ht_0})} \right)
\]

(1)

where the second equality uses the fact that \( e_{ht}(p_{ht}, v_h(p_{ht}, I_{ht}; \zeta_{ht_0}); \zeta_{ht_0}) = I_{ht} \) and where \( u_{ht} \equiv v_h(p_{ht}, I_{ht}; \zeta_{ht}) \) represents utility achieved under the time \( t \) budget set and preferences. Welfare changes equal the change in household nominal income deflated by the change in the expenditure function in response to changes in prices, evaluated along the initial indifference curve. The deflator is evaluated at the initial indifference curve because, by the definition of \( CV_h \), the income compensation it receives at \( t_1 \) leaves the household on that indifference curve.

To understand what one needs to know in order to construct the price deflator in (1), consider any smooth path of prices from \( p_{ht_0} \) to \( p_{ht_1} \), where \( t \) indexes “time” (or, more generally, increments along which prices change between two points \( t_0 \) and \( t_1 \)). Using Shephard’s lemma, (1) can be expressed as (see Lemma 1 in Baqae and Burstein, 2023)

\[
CV_h = \log \left( \frac{I_{ht_1}}{I_{ht_0}} \right) - \int_{t_0}^{t_1} \sum_i b_{hi}^{CV}(p_h) \frac{d}{dt} \log p_{ih} dt
\]

(2)

where \( b_{hi}^{CV}(p_h) \equiv b_{hi}(p_h, u_{ht_0}; \zeta_{ht_0}) \) represents household \( h \)’s budget share on good \( i \) at prices \( p_h \) along its initial indifference curve. Equation (2) implies that welfare changes for a consumer with non-homothetic preferences that are subject to taste shocks are identical to welfare changes for a fictional consumer with homothetic and stable preferences with budget shares as a function of prices given by \( b_{hi}^{CV}(p_h) \).\(^{18}\)

**Discussion.** Equations (1) and (2) hold globally—for changes in prices and incomes of any size. According to equation (2), in order to measure \( CV \) (for given price changes) we only need to know compensated budget shares as a function of prices, \( b_{hi}^{CV}(p_h) \). Given

\(^{18}\)That is, \( b_{hi}^{CV}(p) \) corresponds to the budget shares of a fictional consumer with homothetic preferences represented by the expenditure function \( e_{h}^{CV}(p, u) = e_{h}(p, u_{ht_0}; \zeta_{t_0})u \).
the path of prices from \( t_0 \) to \( t_1 \), these budget shares can be constructed from initial budget shares, \( b_{hit}^{CV}(p_{hit}) \), and cross-price elasticities between all goods along the initial indifference curve. Given these cross-price elasticities, measuring \( CV_h \) does not require income elasticities or taste shifters.\(^{19}\) However, in estimating these cross-price elasticities, shifts in demand induced by income effects and taste shocks cannot be ignored, as we discuss in Section 4.

In our quantification of the welfare impacts of factual or counterfactual changes in prices, we directly measure initial budget shares over consumption goods in our Swiss data. We specify particular preferences to estimate cross-price elasticities along the initial indifference curve using the price changes induced by the 2015 CHF appreciation.

### 3.2 Non-homothetic CES preferences

In what follows, we restrict the general setup of Section 3.1 by imposing non-homothetic CES preferences. There are multiple sectors, indexed by \( s \), and within each sector there is a fixed set of goods, indexed by \( i \in \mathcal{I}(s) \), some imported and some produced domestically.

The expenditure function is given by

\[
e_h(p_{ht}, u; \zeta_{ht}) = f_h(u) \left[ \sum_s \zeta_{hst} u^{\gamma_s} (P_{hst})^{1-\rho} \right]^{\frac{1}{1-\rho}} \tag{3}
\]

\[
P_{hst} = \left( \sum_{i \in \mathcal{I}(s)} \zeta_{hit} u^{\gamma_i} (p_{hit})^{1-\eta_s(u)} \right) \frac{1}{1-\eta_s(u)} \tag{4}
\]

where \( f_h(\cdot) > 0 \) and \( \rho, \eta_s(\cdot) \in [0, 1) \cup (1, \infty) \).\(^{20}\) By Shephard’s lemma, the budget share of any good \( i \in \mathcal{I}(s) \) is

\[
b_{hit} = \frac{\zeta_{hit} u^{\gamma_i} p_{hit}^{1-\eta_s(u_{hit})}}{\sum_{i' \in \mathcal{I}(s)} \zeta_{hi't} u^{\gamma_{i'}} (p_{hi't})^{1-\eta_s(u_{hit})}} \times b_{hst} \tag{5}
\]

\(^{19}\)If we used the equivalent (rather than compensating) variation under final (rather than initial) preferences, then computing welfare changes requires budget shares as a function of prices along the final (rather than initial) indifference curve. Since in our applications we consider the welfare implications for Swiss consumers of counterfactual price changes starting in 2014, it is more convenient to focus on CV (which requires estimates of price elasticities in 2014) rather than EV (which requires budget shares and estimates of price elasticities in an unobserved initial period such as autarky).

\(^{20}\)These preferences reduce to nested homothetic CES if, for example, \( \eta_s(u) \) is independent of \( u \), \( \gamma_i = \gamma_s = 0 \) for all \( i \) and \( s \), and \( f'_h(u) > 0 \). See Appendix C for additional information on these preferences.
where \( b_{hst} \equiv \sum_{i \in I(s)} b_{hit} \) is the share of sector \( s \) in \( h \)'s budget at time \( t \), given by

\[
b_{hst} = \frac{\zeta_{hst} u_{hit}^{1-\rho} P_{hst}^{1-\rho}}{\sum_{s'} \zeta_{hs't} u_{hit}^{1-\rho} P_{hs't}^{1-\rho}}. \quad (6)
\]

As described in detail below, in mapping our model to the data we consider three aggregate sectors \( s \) listed in Table 1. Within each sector we either map each \( i \) to a homothetic aggregator across domestic products and a homothetic aggregator across imported products (we do not introduce notation for these aggregators to simplify presentation) or we map each \( i \) to individual barcode products. In the first approach, \( \eta_s(u) \) is the elasticity of substitution between the aggregate domestic good and the aggregate imported good in sector \( s \). In the second approach, \( \eta_s(u) \) is the elasticity of substitution between any pair of barcode products in sector \( s \), irrespective of import status.

**Welfare changes.** Compensated budget shares \( b_{CV_{hi}}(p_h) \) are obtained by fixing utility at \( u_{hit_0} \) and tastes at \( \zeta_{hit_0} \) and \( \zeta_{hst_0} \). We express compensated budget shares as a simple function of initial expenditure shares, initial elasticities of substitution, and changes in prices:

\[
b_{CV_{hi}}(p_h) = b_{hit_0} \times \left( \frac{\tilde{p}_{hi}}{\tilde{p}_{hs}} \right)^{1-\eta_{hst_0}} \times \frac{\tilde{p}_{hs}^{1-\rho}}{\sum_{s'} b_{hs't_0} \tilde{p}_{hs'}^{1-\rho}}. \quad (7)
\]

\[
\tilde{p}_{hs} \equiv \left( \sum_{i \in I(s)} \frac{b_{hit_0}}{b_{hst_0}} \left( \frac{\tilde{p}_{hi}}{\tilde{p}_{hs}} \right)^{1-\eta_{hst_0}} \right)^{1-\eta_{hst_0}}. \quad (8)
\]

where \( \tilde{x} \equiv x/x_{t_0} \) for any \( x \); \( \rho \) is the elasticity of substitution along the initial indifference curve between sectors, which we assume is common across all households and constant; and \( \eta_{hst_0} \) is the elasticity of substitution along the initial indifference curve for household \( h \) between goods within sector \( s \).

Given compensated budget shares, the expression for welfare changes in the general setup, (2), simplifies to

\[
CV_h = \log \left( \bar{I}_h \right) - \frac{1}{1-\rho} \log \left[ \sum_s b_{hst_0} \left( \tilde{p}_{hs} \right)^{1-\rho} \right]. \quad (9)
\]

where \( \tilde{p}_{hs} \) is defined by equation (8).

We use (9) to construct changes in welfare in response to factual and counterfactual income and price changes. Constructing \( CV_h \) for household \( h \) requires the value of the elasticity of substitution between sectors, \( \rho \), expenditure shares across sectors in the ini-
tial period, $b_{ht0}$, income changes, $\hat{I}_h$, and sectoral price changes $\hat{P}_s$. Constructing $\hat{P}_s$ in equation (8) for household $h$ requires expenditure shares within sector $s$ in the initial period, $b_{ht0}$, and the elasticity of substitution in the initial period $t_0$, $\eta_{ht0}$.

To a second-order approximation, and setting $\rho \to 1$, equation (9) can be expressed as

$$CV_h \approx \log (\hat{I}_h) - E_{b_{ht0}} \left[ \log \hat{P} \right] + \frac{1}{2} \sum_s b_{ht0} \left( \eta_{ht0} - 1 \right) \text{Var}_{b_{ht0}|s} \left[ \log \hat{P} \right]$$

(10)

where $E_{b_{ht0}} \left[ \log \hat{P} \right]$ is the budget-share weighted average of log price changes and where $\text{Var}_{b_{ht0}|s} \left[ \log \hat{P} \right]$ is the budget-share weighted variance of log price changes within sector $s$; see Baqaee and Burstein (2023). The approximation error in expression (10) vanishes as price changes become smaller and as $\eta_{ht0} \to 1$. The literature on the unequal effects of price changes has largely focused on the first-order effects in equation (10). The expenditure-switching effect, which is the focus of our paper, raises welfare if the elasticity of substitution $\eta_{ht0}$ is greater than one, and is increasing in $\eta_{ht0}$. That is, households that are more price sensitive benefit more from volatility in prices (or lose less if $\eta_{ht0} < 1$).

**Unequal expenditure switching.** Here, we provide two special cases that highlight the importance of differences in elasticities, $\eta_{ht'0} - \eta_{ht0}$, for differences in $CV_{ht'}$ and $CV_h$. In both cases, we set $\rho \to 1$.

First, consider two households with common expenditure shares in the initial period and arbitrary changes in product prices. To a second-order approximation, the difference in the expenditure switching effect between these households is

$$\frac{1}{2} \sum_s b_{ht0} \left( \eta_{ht'0} - \eta_{ht0} \right) \text{Var}_{b_{ht0}|s} \left[ \log \hat{P} \right]$$

(11)

which depends on the difference in their elasticities of substitution in the initial period, $\eta_{ht'0} - \eta_{ht0}$. Conditional on this difference, the levels of these elasticities do not matter to a second-order approximation. It is precisely this difference in elasticities that we estimate in Section 4.\(^{21}\)

Second, consider two households with (potentially) different expenditure shares within

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\(^{21}\)Allowing for differences in initial budget shares, the differences in the expenditure-switching effect includes the additional term

$$\frac{1}{2} \sum_s \left( \eta_{ht'0} - \eta_{ht0} \right) \left( b_{ht'0} \text{Var}_{b_{ht'0}|s} \left[ \log \hat{P} \right] - b_{ht0} \text{Var}_{b_{ht0}|s} \left[ \log \hat{P} \right] \right)$$

This additional term depends on the level of the elasticity. However, since the differences in price variances across households is small in our quantitative application, this additional term is small.
sectors and assume that the distribution of log price changes, $\log \hat{p}_{hitr}$, in each sector is normal with mean $\mu_s$ and standard deviation $\sigma_s$. In this case, equation (10) is exact. Moreover, if expenditure shares across sectors are common across households, then the difference in the expenditure switching effect between them is given by equation (11), where $\text{Var}_{b_{ht0}|s} \left[ \log \hat{P} \right]$ is simply $\sigma^2_s$. Conditional on differences in $\eta_{ht0} - \eta_{hst0}$ the levels of these elasticities do not matter, now globally.

**Discussion of preferences.** The non-homothetic CES preferences we consider are general in a number of ways. First, they allow for non-unitary income elasticities that vary across goods within sectors as can be seen in equation (5), driven by differences in $\gamma_i$ across goods and the dependence of price elasticities on $u$, and also across sectors as can be seen in equation (6). As shown in equation (2) and discussed in Section 3.1 in the general formulation, income elasticities play no role in the construction of the CV conditional on knowing initial expenditure shares and compensated cross-price elasticities. Second, these preferences allow for elasticities of substitution that vary across households as a function of utility $u_h$, as in Fally (2022). As shown in equation (9), calculating $CV_h$ requires values for these elasticities of substitution in the initial period.

Contrary to this generality, these preferences impose strong restrictions. Elasticities of substitution are constant along any indifference curve as in standard CES models. We make this assumption for three reasons. First, we estimate these elasticities of substitution leveraging the 2015 Swiss franc appreciation, which does not contain sufficient price variation to estimate them globally. Second, this restriction has an appealing theoretical property: it implies that the integral over prices in equation (2) simplifies substantially, as shown in equation (9). It additionally implies that compensated budget shares in equations (7) and (8) and the CV in equation (9) for a particular household are identical to those in a model in which the household has homothetic and stable CES preferences with household-specific, exogenously given, and constant demand shifters and elasticities.

Third, this restriction implies that only a small subset of preference parameters are required for measuring $CV$, as opposed to other demand systems, e.g. the Almost Ideal Demand System, in which the same parameters control both income and cross-price elas-

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22Fally (2022) establishes sufficient conditions for the rationalization of non-homothetic CES demand when the elasticity of substitution is a decreasing function of $u$, which is the empirically relevant case in our data. In Appendix C we show—under certain assumptions—that these conditions are satisfied under the parameterization of $\eta(u)$ that we assume to derive our main estimating equation. We also describe numerical simulations for which the expenditure function is monotonically increasing in $u$ for a wide range of parameters.

23It is standard to calculate changes in price indices across households imposing homothetic CES preferences with demand shifters and elasticities that vary across households but are fixed in the counterfactuals (see, e.g., Handbury 2021). Our results show that this is equivalent to calculating changes in the welfare-relevant deflator when preferences are generalized non-homothetic CES.
ticities.

The other strong restriction we impose is that for any household there is a single elasticity, $\rho$, that shapes substitution between sectors and a single elasticity, $\eta_{hst}$, that shapes patterns of substitution between goods within sector $s$. This dramatically reduces the dimensionality of the problem. This formulation is equivalent to one with additional nests (e.g. product categories) under the assumption that the elasticity of substitution within nests is equal to the one between nests.

4 Elasticities of substitution and income

In this section we estimate differences in compensated price elasticities across incomes using the Homescan data, where we observe household-product-specific expenditure shares and prices.

4.1 Estimating equation

To estimate how elasticities of substitution vary with income, we must take into account that changes over time in budget shares reflect not only price changes but also income effects and demand shifters. For any continuing good, differentiating equation (5) at $t_0$ yields

$$
d \log b_{hit} = d \log \zeta_{hit} + \left( \gamma_i - \frac{\partial \eta_s}{\partial u_h} u_{hit0} \log p_{hit0} \right) d \log u_{hit} + (1 - \eta_{hst}) d \log p_{hit} + \psi_{hst}
$$

(12)

where $\psi_{hst} \equiv d \log \left( \frac{b_{hst}}{\sum_{i' \in \mathcal{I}(s)} \zeta_{hi} u_{hi} \eta_i \eta_s} \right)$ and all derivatives (in the previous and subsequent equations) are evaluated at $t_0$. Differentiating $I_{ht} = e_h (p_{ht}, u; \zeta_{ht})$,

$$
d \log u_{ht} = \left( \frac{\partial \log e_h}{\partial \log u_h} \right)^{-1} \times \left( d \log \frac{I_{ht}}{P_{ht}} - \bar{\epsilon}_{ht} \right)
$$

(13)

where $d \log (I_{ht}/P_{ht})$ is the change in income deflated by a household-specific weighted average of price changes across goods in all sectors $d \log P_{ht} \equiv \sum_i b_{hit0} d \log p_{hit}$, and $\bar{\epsilon}_{ht} \equiv \sum_i \left( \frac{d \log e_h}{\partial \zeta_{hit}} \right) d \zeta_{hit}$ is the shift in the expenditure function due to taste shifters; see Appendix C for derivations. We refer to $d \log (I_{ht}/P_{ht})$ as the change in real income for
To estimate how elasticities of substitution vary with initial income, we impose two restrictions (in Appendix C we provide a cardinalization of the utility function that microfound these two restrictions). These restrictions play no role for our counterfactual welfare calculations conditional on estimates of elasticities of substitution; we impose these restrictions to facilitate the estimation of these elasticities. First, we assume that household h’s income elasticity for good i at \( t_0 \), \( \frac{\partial \log b_{hit}}{\partial \log I_{ht}} \), can be expressed as the sum of a good i-specific term that is common for all households, which we denote by \( \kappa_i \), and a household-sector-specific term. This assumption holds if the term multiplying the change in real income in expression (14) evaluated at \( t_0 \) can be written as

\[
\left( \frac{\partial \log e_h}{\partial \log u_h} \right)^{-1} \times \left( \gamma_i - \frac{\partial \eta_s}{\partial \log u_h} \log p_{hit0} \right) = \kappa_i + \kappa_{hs} \tag{15}
\]

Second, we assume a log-linear relation between the elasticity of substitution in sector s and household income in the initial period,

\[
\eta_{hst_0} \equiv \bar{\eta}_s + \eta_s \log I_{ht0}. \tag{16}
\]

If \( \eta_s < 0 \), then a higher-income household is less price sensitive in sector s at \( t_0 \).

Under these two additional restrictions, equation (14) can be expressed as

\[
d \log b_{hit} = v_{hit} + \kappa_id \log \left( \frac{I_{ht}}{P_{ht}} \right) + [1 - \bar{\eta}_s - \eta_s \log(I_{ht0})]d \log p_{hit} + \tilde{\psi}_{hst}. \tag{17}
\]

The first term, \( v_{hit} \equiv d \log \tilde{\xi}_{hit} - \kappa \tilde{\epsilon}_{ht} \), corresponds to household h’s demand shifter for good i due to taste shocks. The second term captures the interaction between the good i-specific component of the income elasticity and the change in real income for household h, giving rise to a demand shifter for good i due to income effects. The third term corresponds to the compensated price elasticity for good i in the initial period interacted with the change in the price of good i. The last term, \( \tilde{\psi}_{hst} \), groups all factors that vary at the sector × household level.\(^{24}\)

\(^{24}\)Setting changes in income, tastes, and prices of goods \( j \neq i \) equal to zero, equation (17) resembles the familiar Slutsky equation relating Marshallian, Hicksian, and income elasticities. Our baseline approach to
We can decompose the demand shifter $v_{hit}$—without loss of generality—into the component of the demand shock for good $i$ that is common across all households, a demand shock for imports that varies freely across households, and a household-good-specific deviation from these. Specifically, $v_{hit} \equiv v_{it} + \tilde{FE}_{hst}I_{im}^M + \tilde{v}_{hit}$, where $I_{im}^M$ is an indicator variable that equals one if good $i$ is imported. This yields our baseline estimating equation

$$d \log b_{hit} = FE_{it} + FE_{hst}^M + \kappa_i d \log \left( \frac{I_{ht}^M}{p_{hit}} \right) - \eta_s \log(I_{ht0}) d \log p_{hit} + \iota_{hit}.$$ (18)

In equation (18), the product-fixed effect $FE_{it}$ is the sum of the average product-specific demand shock across households, $v_{it}$, and the common impact of the average price change for product $i$ across households, $(1 - \bar{\eta}_s)d \log p_{it}$; the term $FE_{hst}^M \equiv \psi_{hst} + \tilde{FE}_{hst}I_{im}^M$ is a household $\times$ import status fixed effect; and finally, the term $\iota_{hit} \equiv \tilde{v}_{hit} + (1 - \bar{\eta}_s)(d \log p_{hit} - d \log p_{it})$ is a residual that includes both the household’s demand-shock deviation for product $i$ (relative to the average across households and, if $i$ is imported, the household’s average demand shock for imported goods) as well as the common effect of the household-specific deviation in the change in product $i$’s price relative to its average change across households.

We estimate variation in elasticities of substitution across households, $\eta_s$, using equation (18) in two different ways leveraging distinct sources of variation. In our first approach, we use variation in changes in import relative to domestic expenditures across higher- and lower-income households, similar to the variation in Stylized Fact 3. In our second approach, we use variation in changes in expenditures across individual barcode products and variation in product price changes across aggregations of higher- and lower-income households. The advantages of the first approach are simplicity, the ability to estimate equation (18) at the household level, and the straightforward connection to Stylized Fact 3. We view this as a first pass. The benefit of the second approach is that it substantially relaxes our identification assumption: it is valid in the presence of entry and exit of products and in the presence of import demand shocks that vary systematically with income (which we cannot a priori rule out, even though the 2015 CHF appreciation was triggered by a policy response to foreign events and took place in the context of a stable Swiss economy both in terms of aggregates and nominal income inequality). In both approaches, we estimate equation (18) taking differences between 2014 and 2015. Even though these two approaches leverage entirely distinct sources of variation to identify $\eta_s$, they yield remarkably similar results.

estimating differences in Hicksian price elasticities does not require estimating income elasticities for each good using cross-sectional data. In Appendix B.2 we consider an alternative procedure that relaxes restriction (15) but requires first estimating income elasticities in the cross section under additional assumptions.
4.2 Approach 1: Import and domestic expenditures by household

In our first approach, we assume that there are only two goods within groceries: an imported good $i = M$ and a domestic good $i = D$.\footnote{To map this first approach to our data (in which there are multiple imported and domestic goods), we assume that each of the two goods is itself a stable homothetic aggregator across a fixed set of imported and domestic varieties. In constructing import and domestic prices within groceries to use in the estimation, we use a first-order approximation of the expenditure function of any homothetic aggregator within each import status.} In this case, whereas we can control for an aggregate import demand shock (contained in $\mathbb{F}E_{it}$ in equation 18), we cannot control for a household-specific import demand shock, since there is only one imported good. Hence, $\mathbb{F}E_{hst}^M$ reduces to a household effect. Since there are only two goods, we take differences across the imported and domestic goods and, since there are only two time periods, we also replace the time effect with a constant and estimate

$$d \log \left( \frac{b_{hMt}}{b_{hDt}} \right) = \alpha + \kappa d \log \left( \frac{I_{ht}}{P_{ht}} \right) - \eta_s \log(I_{ht})d \log \left( \frac{p_{Mt}}{p_{Dt}} \right) + \iota_{ht} \tag{19}$$

Here, $\alpha$, $\kappa$, and $\iota_{ht}$ all represent differences of the parameters in equation (18) across imported and domestic goods: $\alpha \equiv \mathbb{F}E_{M} - \mathbb{F}E_{D}$, $\kappa \equiv \kappa_M - \kappa_D$, and $\iota_{ht} \equiv \iota_{hM} - \iota_{hD}$. We measure $b_{hDt}$ and $b_{hMt}$ as the expenditure shares on domestic and imported goods within each individual household. The price changes for imported and domestic goods are measured as weighted averages of annual changes in national prices of products weighted by expenditures per product across all consumers in 2014, separately for imports and domestic goods. By using a single national relative import price change, our estimating equation in approach 1 is very similar to that used in columns 1–3 of Table 2, where we documented Stylized Fact 3. We measure household $h$’s inflation rate (across all sectors), $d \log P_{ht}$, using disaggregated price data in the CPI as measured by the SFSO (these price changes are common across households) and income group-specific expenditure shares across these disaggregated categories. We measure annual changes in nominal income by household using a Swiss household panel on income (FORS); details are available in Appendix A.

Identification. We identify differences in elasticities across household incomes from changes in import expenditure shares that are correlated with household income. We do not instrument for price changes since we have only one value of $d \log (p_{Mt}/p_{Dt})$. The identification assumption estimating regression (19) using OLS is that household-specific import demand shock deviations from the aggregate import demand shock between 2014 and 2015 are uncorrelated across household incomes in 2014. We relax this restriction in Ap-
Table 4: Estimation of $\eta_s$ in Approach 1 using equation (19)

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
<th>(6)</th>
<th>(7)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\log(I_{ht_0})d\log(P_{Mt}/P_{Dt})$</td>
<td>2.189***</td>
<td>2.207***</td>
<td>1.838***</td>
<td>1.981***</td>
<td>2.041***</td>
<td>2.172***</td>
<td>2.361***</td>
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<td></td>
<td>[0.554]</td>
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<td>0.557***</td>
<td>0.469***</td>
<td>0.492***</td>
<td>0.516***</td>
<td>0.550***</td>
<td>0.581***</td>
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<td>[0.132]</td>
<td>[0.110]</td>
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<td>X</td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td>X</td>
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</tr>
</tbody>
</table>

Notes: The estimating equation is (19). We report $-\eta_s$ and $\alpha$. Observations are households and the dependent variable is the log change in import-relative-to-domestic expenditures across all Homescan products between 2014 and 2015. In our baseline in column 1, robust standard errors are clustered by 50 household bins according to household income in 2014, observations are weighted by the product of the number of households in each of the 50 bins and the share of each household’s expenditures among households within that bin, and we winsorize the dependent variable at the first percentile in both tails. Columns 2-7 each make one change relative to the baseline in column 1. In column 2 we do not winsorize, in column 3 we instead winsorize at the 5th percentile, in column 4 we do not weigh observations, in column 5 instead weigh observations by expenditures in 2014, in column 6 we omit changes in real income from the regression specification, and in column 7 we control for household size. *p<.1; **p<.05; ***p<.01

Results. Whereas we estimate regression (19) at the household level, we cluster standard errors by 50 income bins defined by quantiles of the household income distribution in our sample in 2014. We do so to allow for the possibility of correlated imported demand shocks across households in the same income bin; however, as stated above, we continue to require that import demand shocks between 2014 and 2015 across income bins are not systematically related to income. We weigh observations (households) by the product of the number of households in each of the 50 bins and the share of household $h$’s expenditures among households within that bin and winsorize the dependent variable at the first percentile in both tails. We revisit each of these choices in robustness.

The first column of Table 4 displays our baseline results. In all tables we report the estimated coefficient, which is $-\eta_s$. We find $\eta_s = -2.19$, which implies a substantially lower elasticity of substitution for higher-income households. For example, the elasticity of substitution of a household with 2014 income of 60,000 CHF is approximately 2.4 ($\approx 2.19 \times \log 3$) lower than a household with income of 20,000 CHF. This gap shapes the non-linear effects of import price changes in our quantification.
4.3 Approach 2: Product-level expenditures by income group

In our second approach, each $i$ is an individual barcode product. Given the granularity of this definition of a product $i$ relative to our first approach, the household-product-level data are sparse. Hence, to estimate equation (18) we aggregate product-level data across groups of households, as is standard in demand estimation. Specifically, we group households into 50 income bins defined by quantiles of the household income distribution in our sample in 2014; $h \in \{1, ..., 50\}$ now denotes the income bin. Within each bin, we take the median value of 2014 income and the median annual change in nominal income between 2014 and 2015 across individual households. We measure inflation and changes in income at the household level as described above.

In our baseline we use a common price change across households at the product level, $d \log p_{hit} = d \log p_{it}$. In robustness we consider a more disaggregated household aggregation that incorporates spatial variation; in this case, we measure a common $d \log p_{hit}$ for all households within a one-digit zip code and allow these price changes to vary across space.

Identification. In our second approach, we identify differences in elasticities across households from differences across the income distribution in the relationship between changes in expenditure shares and prices at the barcode product level. In this case, we can explicitly incorporate import demand shocks that are specific to each of the fifty household aggregates.

There are two remaining endogeneity concerns. The first is measurement error, which generates attenuation bias. The second is an economic argument for endogeneity. Suppose that high-income households are less price sensitive, consistent with our findings in Approach 1. Consider a product that faces demand shocks that are higher for higher-income households (we control for the average demand shock across households, so only deviations from the average remain in the residual). In response, the firm will face a more inelastic demand and will, therefore, charge a higher markup. Hence, there is a positive correlation between household-specific demand shock deviations (i.e., the residual) and the interaction between product-specific price changes and household initial income (the independent variable of interest). This implies that under the hypothesis that high-income households are less price sensitive, OLS is upward biased.

We address these concerns by constructing an instrument using an interaction between a product-specific cost shifter and initial household income. Our cost shifter ex-

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26Our baseline sample is the set of products that were purchased at least once per month (nationally) in the year-and-a-half before and after the CHF appreciation.
exploits variation across imported goods in the invoicing currency of prices at the border. As described in Section 2.1, we match products to border groups and measure the share of imported products in each border group that are denominated in EUR (out of those denominated in either EUR or CHF) in 2014, which we denote by \( \text{share}_{it0} \). Because of the stickiness of import prices at the border in their invoicing currency, Swiss retail prices of imported goods are more responsive to the CHF appreciation if imports are denominated in EUR rather than in CHF; see Auer et al. (2021).

Since the expected reduction in Swiss retail prices in response to the CHF appreciation is greater for imported products that belong to border groups with a higher fraction of border prices invoiced in EUR, we construct our instrument as the interaction between (i) the share of imported goods in the corresponding border group that are denominated in EUR, \( \text{share}_{it0} \), (ii) an import indicator variable, \( I_{i}^{M} \), and (iii) the logarithm of initial household income, \( \log(I_{ht0}) \). If we restrict our sample to imported goods, as we do in robustness, then the instrument is the interaction between (i) and (iii) alone: \( \log(I_{ht0}) \cdot \text{share}_{it0} \).

In this case, we leverage the fact that the expected reduction in Swiss retail prices among imported goods in response to the CHF appreciation is greater for those goods belonging to border groups with a higher fraction of border prices invoiced in EUR.\(^{27}\)

In some border groups, the number of border price observations denominated in either EUR or CHF with which to construct (i)—the share of imported goods that are denominated in EUR (out of those denominated in EUR or CHF)—is small and, therefore, the share is unreliable. Hence, in our baseline we restrict the sample of products to those in border groups with more than 28 border price observations in 2014 and vary this cutoff in robustness. This leaves 35 border groups in our baseline estimation sample, 7 of which have no imported goods denominated in EUR (i.e., \( \text{share}_{it0} = 0 \) for all products \( i \) in these 7 border groups).

**Results.** In our baseline, we weigh observations by the product of the number of households in each aggregation \( h \) and the share of expenditures among households within that aggregation on product \( i \).\(^{28}\) In constructing changes in log expenditure shares, we winsorize changes in log expenditures at the first percentile (both in the right and left tails). Finally, we two-way cluster standard errors at the level of household income (there are 50

\[^{27}\]There is a large literature (see, e.g., Gopinath and Itskhoki 2022) arguing that the data on invoicing currency of export and import prices are consistent with models in which firms’ invoicing currency choices are based on desired pass-through to exchange rate movements. A sufficient condition for our exclusion restriction is that heterogeneity in anticipated relative demand shocks across the income distribution between 2014 and 2015 does not shape pre-shock invoicing currency choices.

\[^{28}\]This approach puts equal weight on each underlying household rather than giving a higher weight to those household aggregations with higher expenditures (since our objective is to estimate how price sensitivities vary with income).
Table 5: Estimation of $\eta_s$ in Approach 2 using equation (18)

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>OLS</td>
<td>RF</td>
<td>2SLS</td>
</tr>
<tr>
<td>$\log(I_{ht0}) \times d \log p_{it}$</td>
<td>0.018</td>
<td>1.930**</td>
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</tr>
<tr>
<td></td>
<td>[0.134]</td>
<td>[0.867]</td>
<td></td>
</tr>
<tr>
<td>$\log(I_{ht0}) \times \text{share}_{it} \times I_i^M$</td>
<td>-0.140**</td>
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<td></td>
</tr>
<tr>
<td></td>
<td>[0.068]</td>
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</tr>
<tr>
<td>Observations</td>
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<td>95,325</td>
<td>95,325</td>
</tr>
<tr>
<td>K-P F Stat (first stage)</td>
<td>13.1</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: The estimating equation is (18). Observations are barcode product $\times$ household aggregates, where households are aggregated into 50 bins according to initial income. The dependent variable is the log change in expenditures between 2014 and 2015. Column 1 reports OLS results, column 2 reports reduced-form results in which we replace $\log(I_{ht0})d \log p_{it}$ with $\log(I_{ht0})\text{share}_{it}I_i^M$, and column 3 reports 2SLS results in which we instrument for $\log(I_{ht0})d \log p_{it}$ with $\log(I_{ht0})\text{share}_{it}I_i^M$. Robust standard errors are two-way clustered at the level of household income bin and, separately, the interaction between import status and the share of imported goods denominated in EUR in the border group; observations are weighted by the product of the number of households in each aggregation and the share of expenditures among households within that aggregation on product $i$; and we winsorize changes in log expenditures at the first percentile (both in the right and left tails). *p<.1; **p<.05; ***p<.01

such clusters) and, separately, the interaction between import status and the value of the share of imported goods denominated in EUR in the corresponding border group (there are 54 such clusters). We revisit each of these choices in robustness.

Table 5 displays our baseline results, focusing on the parameter of interest: $\eta_s$. The first column reports results from estimating equation (18) using OLS, where we find an economically small and statistically insignificant estimate. Column 2 reports results from estimating the reduced-form specification, in which we replace the change in product price interacted with the logarithm of initial household income with the instrument. We find that, between 2014 and 2015, higher-income households increased their expenditures by less on imported goods within border groups with a higher share of products invoiced in EUR (those with a larger decline in border and retail prices in response to the 2015 CHF appreciation) conditional on real income changes, import demand shocks that vary freely across household income groups, and other covariates. This is the expected sign of the reduced-form relationship.

Column 3 reports the baseline version of our main empirical result, the two-stage least squares estimate of $\eta_s$. The first-stage coefficient is $-0.073$ (implying that, on average, the price of an imported product in a border group that is entirely invoiced in EUR fell by 7.3% more than a product in a border group entirely invoiced in CHF in response to the roughly 14% appreciation of the CHF) and the associated F statistic is 13.1. The second-stage coefficient of interest, $\eta_s = -1.93$, is very similar to the estimate in our first approach, which leverages an entirely distinct source of identification. The bias of the

29Throughout, we report the Kleibergen-Paap Wald rk F statistic when there is only one endogenous variable.
OLS estimate is as expected both due to measurement error and the economic argument described above (recall that we display the negative value of the OLS coefficient, since this is the structural parameter of interest).

4.4 Robustness and sensitivity

Table 4 displays robustness across a range of choices in our first approach. In column 2 we do not winsorize the dependent variable; in column 3 we instead winsorize at the fifth percentile. In column 4 we do not weigh observations and in column 5 we instead weigh by household expenditure in 2014. In column 6 we omit changes in real income from the estimating equation. Finally, in column 7 we control for household size in case it is correlated with income and elasticities vary with household size. Results are largely robust to these choices.

The majority of our robustness exercises focus on our second approach. Most of these—such as incorporating spatial variation into the estimation, varying choices in the construction of household income, and determining whether particular income groups drive our results—are contained in Appendix B.2. Here, we describe a small number of sensitivity and robustness exercises, with results displayed in Table 6.

Recall from Table 2 in Section 2.2 that the gap between the import shares of low- and high-income households fell between 2013–14 and rose both between 2014–2015 and 2015–2016. Changes between 2013–14 might suggest pre-existing trends that would call into question our baseline results. Changes between 2015–16 might suggest mechanisms generating lags in expenditure-switching responses.

Here, we begin by showing that there are no such pre-trends in Approach 2. Column 1 of Table 6 replicates our baseline reduced-form specification. Column 2 of Table 6 documents an absence of pre-existing trends in the reduced-form specification; we cannot study pre-trends in the structural specification since our instrument has no power before the CHF appreciation, given that the CHF-EUR exchange rate is stable between 2013–14. Whereas in our baseline we obtain a coefficient of −0.14 that is significant at the 5% level, running the same regression but replacing changes in expenditure shares between 2014–15 with changes between 2013–14 yields a coefficient that is three orders of magnitude smaller and statistically insignificant; see column 2 of Table 6. We interpret this lack of differential pre-existing trends as strengthening the structural interpretation.

In additional sensitivity on income effects in Appendix B.2, mentioned in footnote 24, we use an alternative approach that relaxes restriction (15), but requires estimating cross-sectional income elasticities under strong assumptions. We apply this in Approach 1 and show that results are robust in Table A8.

30
Table 6: Robustness of Approach 2

<table>
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<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
<th>(6)</th>
<th>(7)</th>
<th>(8)</th>
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</thead>
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<tr>
<td>( \log(I_{ht0}) \times share_{it0} \times I_{M_i} )</td>
<td>(-0.14^{**})</td>
<td>(-0.00)</td>
<td>(-0.10)</td>
<td>(-0.14^*)</td>
<td>([0.07])</td>
<td>([0.07])</td>
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<tr>
<td>( \log(I_{ht0}) \times d \log p_{it} )</td>
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<td></td>
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<td>(1.93^{**})</td>
<td>(1.93^{**})</td>
<td>(2.19^{**})</td>
<td>(2.15^{**})</td>
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<td>78,800</td>
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<td>X</td>
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<tr>
<td>Additional controls II</td>
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<td>F Stat (first stage)</td>
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<td>15.6</td>
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</tbody>
</table>

Notes: Column 1 replicates our baseline RF specification shown in column 2 of Table 5. Columns 2 and 3 report the same specification, but in which the outcome variable is defined over the period 2013–14 (column 2) and 2015–16 (column 3). Column 4 displays estimates of the RF specification on a sample restricted to imported goods alone (so that \( I_{M_i} = 1 \) for all observations). Column 5 replicates our baseline 2SLS specification shown in column 3 of Table 5. Column 6 displays estimates of the 2SLS specification on a sample restricted to imported goods alone. Column 7 incorporates two additional controls interacted with year: the 2014 import share as well as the 2014 expenditure share on each border group. Column 8 additionally incorporates one more control interacted with year: the 2014 average price of each individual product. In columns 5–8 we report the KP F statistic. *p<.1; **p<.05; ***p<.01

Another concern is that the share of imported goods in each border group that are denominated in EUR is correlated with some other product characteristic and that this other product characteristic is driving the differential patterns of substitution for higher- and lower-income households. Here we show that controlling for additional triple interactions in which we replace the share of imported goods in the border group that are denominated in EUR with other border group or product characteristics (the 2014 import share of each border group, the 2014 expenditure share of each border group, and the 2014 average price of each individual product) does not substantially change our results. Column 5 of Table 6 replicates our baseline 2SLS estimate from column 3 of Table 5. Columns 7 and 8 of Table 6 show that including these additional controls has little effect on results.

Finally, in our baseline, we include both imported and domestically produced goods in our estimation sample. However, the instrumented change in price is common for all domestic goods. Columns 4 and 6 display reduced-form and 2SLS results when we restrict the sample to imports.\(^{32}\) Results are largely unchanged, although we only have

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\(^{31}\)Column 3 similarly replicates our baseline reduced-form specification using changes in expenditure shares between 2015–16. As in the 2013–14 period, we find statistically insignificant results. However, the estimated coefficient is closer to the baseline value. This could represent evidence of dynamics in expenditure-switching responses, which we do not explore further.

\(^{32}\)Restricting the sample to imported goods requires replacing our household-import indicator fixed effect with a household fixed effect, since there is no need (or possibility) to control for differential import demand shocks across income groups. Moreover, our instrument reduces to the interaction between the log of household income and the share of imported goods denominated in EUR in the border group.
26 clusters in one dimension.

4.5 Estimating $\bar{\eta}_s$

Neither of the two approaches in Section 4 identify the intercept $\bar{\eta}_s$ defined in equation (16). However, under stronger assumptions they can be adjusted to do so; see Appendix B.3 for details.

In Approach 1, we can identify $\bar{\eta}_s$ if we assume that the average import demand shifter $\nu_{it}$ is zero between 2014 and 2015. This yields $\bar{\eta}_s \approx 26.6$. In Approach 2, we can identify $\bar{\eta}_s$ if we do not control for the average product-specific demand shock $\nu_{it}$ and, instead, move it to the residual. This yields $\bar{\eta}_s = 20.87$. The first (second) approach implies that the elasticity of substitution is 4.92 (1.76) for a household with income of 20,000 CHF and that this elasticity remains positive for all household incomes below approximately 190,000 (50,000) CHF.

To sum up, we impose weaker assumptions estimating $\eta_s$ than $\bar{\eta}_s$ and our estimates of $\eta_s$ are much more similar across approaches 1 and 2 than are our estimates of $\bar{\eta}_s$. For these reasons, in our quantitative analyses in Section 5, we present results for a range of $\bar{\eta}_s$ and show that differences in welfare across incomes do not depend crucially on these values within a broad range, consistent with our analytic results in Section 3.2.

5 Quantification

In this section, we use our estimates in Section 4 to assess the role of heterogeneous expenditure switching in shaping the welfare implications—measured using equation (9)—of factual and counterfactual changes in prices. In Section 5.1 we quantify changes in the welfare-relevant price index for groceries using observed price changes in the Homescan data. In Section 5.2 we consider changes in welfare in response to counterfactual changes in import relative to domestic prices across all consumer goods.

We make the following choices in both sections. The initial period, $t_0$, is 2014. Differences in within-grocery elasticities of substitution across incomes, $\eta_s = -2$, match our Homescan-based grocery estimates in Section 4. We consider households at three income levels within the range of incomes in our sample—20,000, 60,000, and 120,000 CHF—and choose $\bar{\eta}_s$ so that the lowest elasticity of substitution across the income groups that we consider (at 120,000 CHF) is equal to 3.
Table 7: Welfare-relevant grocery price changes 2014–15

<table>
<thead>
<tr>
<th>Income</th>
<th>Heterogeneous elasticities</th>
<th>Homogeneous elasticities</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td></td>
<td>Elasticity</td>
<td>1st-order Switching</td>
</tr>
<tr>
<td>1: 20,000</td>
<td>6.6</td>
<td>-1.1</td>
</tr>
<tr>
<td>2: 60,000</td>
<td>4.4</td>
<td>-1.2</td>
</tr>
<tr>
<td>3: 120,000</td>
<td>3.0</td>
<td>-1.3</td>
</tr>
</tbody>
</table>

Notes: This table displays changes in the welfare-relevant price index, $P_{hs}$ in equation (8), for the grocery sector in response to observed price changes of individual products in the Homescan data between 2014 and 2015. Rows 1–3 display results for households with incomes of 20,000 CHF, 60,000 CHF, and 120,000 CHF. Columns 1–3 use heterogeneous elasticities (6.6, 4.4, and 3) whereas columns 4–6 impose common elasticities (all set to 6.6). Columns 1 and 4 display the first-order effects, columns 2 and 5 display the second-order effect, and columns 3 and 6 display the exact change.

5.1 Heterogeneous effects of observed Homescan price changes

In this section we quantify changes in the welfare-relevant price index, $P_{hs}$ in equation (8), for the grocery sector in response to observed price changes of individual products in the Homescan data between 2014 and 2015. We do not include non-groceries because we only observe price changes at the CPI level, which is much more aggregated than in the Homescan data, where we observe price changes at the barcode level. We measure these barcode-specific national price changes as described in Section 4.3.33

To calculate $P_{hs}$ in equation (8), we need expenditure shares and elasticities of substitution by income group in 2014. To measure expenditure shares, we divide households into three equal-sized bins based on 2014 income: the lowest, middle, and top bins contain households with 2014 annual income of 20,000 CHF, 60,000 CHF, and 120,000 CHF. For each bin, we calculate the expenditure share on each product and assign this expenditure share to the corresponding household; in sensitivity analysis we show that our main results are robust to using common expenditure shares across all households. The elasticities required to compute the price index are calculated as described above, yielding elasticities of 6.6, 4.4, and 3 for households with incomes of 20,000 CHF, 60,000 CHF, and 120,000 CHF, respectively; in sensitivity analysis we show that our main results are robust to varying $\bar{\eta}_s$ so that the elasticity for households with incomes of 120,000 CHF ranges between 1.5 and 5.

The left-hand panel of Table 7 contains our baseline results with heterogeneous elasticities. Column 1 displays the first-order welfare effect of price changes in groceries (using the second-order approximation in equation 10), which is simply the expenditure-share-weighted average of product price changes within groceries. These range from -1.1

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33To reduce the role of abnormally large price changes on the price index, we drop products with year-to-year price ratios above 3 or below 1/3. This has almost no impact on the 2014–15 results.
percent for households with incomes of 20,000 CHF to -1.3 percent for households with incomes of 120,000 CHF. Column 2 displays the expenditure-switching welfare effect of price changes, which is \((1 - \eta_{hst0})\) times half the expenditure-share-weighted variance of price changes within groceries. Whereas these effects are smaller than the first-order terms (they range from -1.0 percent for households with incomes of 20,000 CHF to -0.3 percent for households with incomes of 120,000 CHF), their variation across incomes is larger. These differences in expenditure switching are driven almost entirely by differences in elasticities; in particular, the weighted variance is very similar across income groups. Column 3 shows the full non-linear effect, which is very similar to the sum of the first-order and expenditure-switching effects. The price change for low-income households is -2.2 percent, which is about 50% larger than the price change for high-income households. This gap is almost identical to the gap in the expenditure-switching effect.

Another way to see the importance of the expenditure-switching effect is to set all elasticities to be equal. In the right-hand panel of Table 7, we display results in which we impose the elasticity of the lowest-income households, 6.6, for all three household groups. The first-order effect, displayed in column 4, is obviously unchanged. However, now the expenditure-switching effect is very similar across households, unlike in column 2; it is not identical across households because of small differences in the expenditure-share-weighted variance of price changes. Price indices across income groups, displayed in column 6, differ by much less than those under heterogeneous elasticities reported in column 3; moreover, these differences are of the opposite sign.

**Sensitivity.** We provide a range of additional results in Appendix D. First, we display welfare-relevant grocery price changes the year before the CHF appreciation (2013–14). The variance of price changes between 2014–15 is one-and-a-half times the variance of price changes between 2013–14, as implied by Stylized Fact 5 above. Hence, the gap between income groups in the expenditure-switching effect is similarly one-and-a-half times larger in 2014–15 than in 2013–14.

We also display results imposing common expenditure shares across households. Whereas the first-order effects are, obviously, now identical across households, the second-order effects are little changed from our baseline. Finally, we also display results for alternative levels of \(\bar{\eta}_s\)—so that the elasticity for households with income of 120,000 CHF ranges between 1.5 and 5—while holding the differences in elasticities across households fixed. Greater substitution generates larger declines in the welfare-relevant price index; however, differences between income groups are not very sensitive even for the large range of \(\bar{\eta}_s\) considered.
5.2 Heterogeneous effects of counterfactual import price changes

In this section we quantify the effect of counterfactual changes in import prices across all consumer goods. Our focus here on import price changes contrasts with our focus in Section 5.1 on observed price changes between 2014 and 2015, which reflect not only the CHF appreciation but also price changes that would have occurred in its absence. Our focus here on all consumer goods contrasts with our focus in Section 5.1 on groceries alone, where import shares do not vary systematically with income and the first-order effects of import price changes are mechanically very similar across households. At the aggregate level, import shares in 2014 are higher among higher-income households, as shown in Stylized Fact 1, yielding heterogeneous first-order effects of import price changes. Finally, to highlight the non-linearities induced by expenditure switching, we consider larger import price shocks like those induced by much larger exchange rate changes (see, e.g., Cravino and Levchenko, 2017) or a movement to autarky.

To model counterfactual price changes, we assume that the price change of any imported \( j = M \) or domestic \( j = D \) product \( i \) in any sector \( s \) is given by

\[
\log \tilde{p}_i = \log \tilde{p}_j + \sigma_j \epsilon_i
\]

where \( \log \tilde{p}_j \) is the uniform component of price changes across all imported or domestic products and \( \sigma_j \epsilon_i \) is the product-level idiosyncratic component of price changes, with \( \epsilon_i \sim \mathcal{N}(0,1) \).

Given our focus on the expenditure-side effects of foreign price shocks, we assume that the log change in income for all households, \( \tilde{I}_h \), is equal to the average change in the log price of domestic goods, \( \tilde{p}_D \), as in single-factor trade models without imported intermediate inputs.\(^34\)

Under these assumptions, changes in welfare in (9) are given by

\[
CV_h = -\frac{1}{1-\rho} \log \left[ \sum_s b_{hs0} \left( \sum_{j \in M,D} \frac{b_{hjst0}}{b_{hs0}} \left( \frac{\tilde{p}_j}{\tilde{p}_D} \sqrt{\frac{\sigma_j^2}{\tilde{p}_D^2}} \right)^{1-\eta_{hs0}} \right)^{1-\eta_{hs0}} \right] \quad (21)
\]

where \( b_{hjst0} / b_{hs0} \) is the share of expenditure in sector \( s \) on either imports \( j = M \) or domestic goods \( j = D \). According to (21), a higher variance \( \sigma_j^2 \) increases \( CV_h \) (for \( \eta_{hs0} > 1 \)), in a similar way that a lower log \( \tilde{p}_j \) does, and this effect is stronger the larger \( \eta_{hs0} \) is. Whereas

\(^34\)In estimating compensated price elasticities, we do not impose this restriction, but instead use actual changes in retail prices by good. In our counterfactuals, if all domestic goods have a common imported intermediate share, then differences between households in welfare changes do not depend on the value of this share, for any given change in import to domestic prices.
we use equation (21) in our analysis below, we gain further intuition in the special case in which \( \sigma_j = \sigma \) and \( \eta_{hst_0} = \eta_{ht_0} \), where changes in welfare are given by

\[
CV_h = -\frac{1}{1-\rho} \log \left[ \sum_s b_{hst_0} \left( \frac{b_{hMst_0}}{b_{hst_0}} \left( \frac{\widehat{p}_M}{\widehat{p}_D} \right)^{1-\eta_{ht_0}} + \frac{b_{hDst_0}}{b_{hst_0}} \right)^{1-\eta_{ht_0}} \right] + \frac{1}{2} \sigma^2 (\eta_{ht_0} - 1) \tag{22}
\]

In this case, changes in welfare are additively separable in the uniform and idiosyncratic components of price changes.

We consider three sectors \( s \): groceries, non-grocery goods, and services. For each of the three income groups (20,000, 60,000, and 120,000 CHF), we construct import shares in each of these aggregate sectors and expenditure shares across them using data on expenditures and import shares within highly disaggregated consumer categories in the SFSO data (see Table 1 in Section 2.1).\(^{35}\) Aggregating up to three sectors has no effect on our measure of overall import shares by income. We impose a value of \( \rho \) very close to one, \( \rho = 0.99 \), and vary this parameter in sensitivity analysis. We choose values for \( \eta_{hst_0} \) as described above.

We quantify the impact of import price shocks, \( \Delta \equiv \log \widehat{p}_M - \log \widehat{p}_D \), for different values of \( \Delta > 0 \) ranging from \( \Delta = 2.2\% \) (the size of the reduction of import prices relative to domestic prices in 2015) to 1,000\% (a movement to autarky, which is a focus of the quantitative trade literature). We begin by imposing \( \sigma_j^2 = 0 \), so that only the uniform component in equation (22) is active.

**Uniform price changes within M and D.** The first panel of Table 8 reports the welfare implications for each household of import price increases of various sizes.

Higher-income groups are harmed more by import price increases for two reasons. First, they have higher import shares, which shape the first-order effect displayed in equation (10). The import shares in 2014 are 21\%, 24\%, and 27\% for households with incomes of 20,000, 60,000, and 120,000 CHF, respectively, as displayed in Table 1. Second, they have lower initial elasticities of substitution, which shape the expenditure-switching effect.

The bottom two panels of Table 8 highlight the increasing importance of the expenditure-switching effect as the size of the change in import prices \( \Delta \) grows. The middle panel displays the percent difference between the \( CV \) of the lowest-income household and the \( CV \) of the middle- and highest-income households. For a 10\% import price increase, the welfare of the middle- and high-income households fall by approximately 22\% and 41\%.

\(^{35}\)In practice, in assigning import shares in each of these three sectors and expenditure shares across them for our household with income of 60,000 CHF, we use an income of 60,252 CHF instead of 60,000 CHF. This is the cutoff separating the first and second income brackets in Table 1.
Table 8: Compensating variation of counterfactual import price shocks

<table>
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<th>Annual income</th>
<th>+2.2</th>
<th>+10</th>
<th>+20</th>
<th>+40</th>
<th>+1000</th>
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<tbody>
<tr>
<td>1: 20,000 elasticity 6.6</td>
<td>-0.4</td>
<td>-1.8</td>
<td>-3.2</td>
<td>-4.7</td>
<td>-5.6</td>
<td>-0.2</td>
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<td>-4.1</td>
<td>-7.0</td>
<td>-11.1</td>
<td>-0.4</td>
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<tr>
<td>3: 120,000 elasticity 3.0</td>
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<td>-2.6</td>
<td>-5.0</td>
<td>-9.1</td>
<td>-22.0</td>
<td>-0.5</td>
</tr>
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% difference in CV btw income groups 2 and 1

<table>
<thead>
<tr>
<th></th>
<th>16</th>
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<th>31</th>
<th>50</th>
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<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>income groups 3 and 1</td>
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<td>41</td>
<td>57</td>
<td>95</td>
<td>295</td>
<td>148</td>
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Contribution of heterogeneous \( \eta \)s

<table>
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<th>44</th>
<th>62</th>
<th>79</th>
<th>69</th>
</tr>
</thead>
<tbody>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>income groups 3 and 1</td>
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<td>25</td>
<td>41</td>
<td>60</td>
<td>86</td>
<td>67</td>
</tr>
</tbody>
</table>

Notes: Percent changes are 100 × the log of the relative price change. “% difference in CV btw income group j and 1” is \((CV_j - CV_1)/CV_1\) for income group \(j\). “Contribution of heterogeneous elasticities” is \(1 - (CV_j^{\text{homog}} - CV_1^{\text{homog}})/(CV_j - CV_1)\) where \(CV_j^{\text{homog}}\) is the compensating variation of income group \(j\) in our alternative counterfactual in which elasticities are common across income groups and set to the value for a household with income of 20,000 CHF. All columns but the last impose \(\sigma^2_j = 0\). In the final column, we set \(\sigma^2_j\) for \(j = D\) and \(j = M\) to match the observed increase in idiosyncratic volatilities between 2013–14 and 2014–15.

respectively, more than for the low-income household. When import prices rise by more, the differences in welfare changes between incomes and the contribution of heterogeneous elasticities to these differences grow substantially.

To quantify the importance of the expenditure-switching effect, we consider an alternative parameterization in which we impose a common price elasticity across incomes equal to that of households with income of 20,000 CHF (which is 6.6). The bottom panel of Table 8 displays the contribution of heterogeneous elasticities (comparing heterogeneous elasticities and import shares to heterogeneous import shares alone) in shaping differences in welfare changes for the middle- and high-income groups compared to the low-income group. Differences in elasticities between the low- and high-income groups explain only 8% (middle vs low income) and 7% (high vs low income) of the difference in welfare changes when the import price rises by 2.2%. However, when the import price rises by 20%, differences in elasticities explain 44% and 41% of the differences in welfare changes. The larger is the increase in import prices, the higher is the contribution of dif-

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36 For income group \(j\), this is simply \(1 - (CV_j^{\text{homog}} - CV_1^{\text{homog}})/(CV_j - CV_1)\) where \(CV_j^{\text{homog}}\) is CV of income group \(j\) in our alternative parameterization with homogeneous elasticities. Another way to quantify the contribution of heterogeneous elasticities is to compare results with heterogeneity in both elasticities and import shares to results with heterogeneity in elasticities alone. These results are very similar to those reported in the bottom panel of Table 8.
ferences in elasticities to the unequal welfare changes across incomes. For a movement to autarky, the expenditure-switching effect accounts for the vast majority (79% and 86%) of the unequal welfare effects.

**Dispersed price changes within M and D.** To this point in Section 5.2, we have considered the uniform component of import price shocks $\Delta \equiv \log \hat{p}_M - \log \hat{p}_D$ imposing a zero variance for idiosyncratic price changes within imported $j = M$ and domestic goods $j = D$. Recall from equation (22) that if $\sigma^2_M = \sigma^2_D$, the welfare impact due to idiosyncratic price changes is additively separable from the uniform component. To evaluate the overall effect, we must calibrate $\sigma^2_j$.

We first consider an import price shock of size $\Delta = 2.2\%$, which is the size of the average decline in import relative to domestic prices observed between 2014–15. Rather than setting $\sigma^2_j$ to the observed variance of price changes between 2014–15 (which includes price changes that would have occurred in the absence of the import price shock), we set it to the observed increase in the variance of price changes between 2013–14 and 2014–15.37

The final column of Table 8 shows that the welfare loss of the lowest-income household is smaller than for the middle- and highest-income households, and more than two-thirds of the differences across incomes is driven by heterogeneity in $\eta$s across households. These results contrast with the first column of Table 8 (no idiosyncratic price changes), where relative differences in welfare across households are smaller and mostly driven by heterogeneous import shares. Intuitively, the price volatility associated with the import price shock reduces its welfare costs for all households, but does so disproportionately for more elastic households.

Finally, we consider a larger import price shock, of size $\Delta = 10\%$. Since the idiosyncratic price volatility generated by such a shock is unobserved, we consider a wide range of volatilities. Each column in Table 9 considers a value of $\sigma^2_j$ that is a factor $x$ of our calibrated variance under the $\Delta = 2.2\%$ shock for various values of $x$. The first column ($x = 0$) corresponds to the second column of Table 8, where we set the idiosyncratic price variance to zero. The second column ($x = 1$) imposes the same idiosyncratic variance as under the smaller, $\Delta = 2.2\%$ shock. As $x$ increases from 0 to 3, differences in wel-

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37To motivate why we set the counterfactual variance equal to the difference in variance between 2013–14 and 2014–15, suppose that the idiosyncratic component of price changes is the sum of a component induced by the import price shock, $\sigma_1 \epsilon_{1i}$, and a component that is orthogonal to the import price shock, $\sigma_2 \epsilon_{2i}$, where $\epsilon_{1i}$ and $\epsilon_{2i}$ are i.i.d. and normally distributed. The sum of the two components can be written as in equation (20), where $\sigma^2_j = \sigma^2_{1j} + \sigma^2_{2j}$. For our counterfactual import price shocks, we set $\sigma_{2j} = 0$. To assign $\sigma_{1j}$ for a 2.2% import price shock, we assume that the variance of price changes between 2013–14 equals $\sigma^2_{2j}$ and the variance of price changes between 2014–15 equals $\sigma^2_{1j} + \sigma^2_{2j}$. Thus, $\sigma^2_{1j}$ equals the variance of price changes between 2014–15 minus the variance of price changes between 2013–14. Specifically, from Table 3, we obtain $\sigma^2_M = 0.0058 - 0.0033 = 0.0025$ and $\sigma^2_D = 0.0023 - 0.0018 = 0.0005$. 

35
Table 9: Compensating variation of a 10% import price shock

<table>
<thead>
<tr>
<th>Annual income</th>
<th>Ratio of variance of idiosyncratic price changes to the calibrated variance with a 2.2% shock</th>
</tr>
</thead>
<tbody>
<tr>
<td>1: 20,000</td>
<td>6.6</td>
</tr>
<tr>
<td>2: 60,000</td>
<td>4.4</td>
</tr>
<tr>
<td>3: 120,000</td>
<td>3.0</td>
</tr>
<tr>
<td>0</td>
<td>-1.8</td>
</tr>
<tr>
<td>1</td>
<td>-1.6</td>
</tr>
<tr>
<td>1.5</td>
<td>-1.5</td>
</tr>
<tr>
<td>2</td>
<td>-1.4</td>
</tr>
<tr>
<td>2.5</td>
<td>-1.3</td>
</tr>
<tr>
<td>3</td>
<td>-1.1</td>
</tr>
<tr>
<td>0</td>
<td>-2.2</td>
</tr>
<tr>
<td>1</td>
<td>-2.1</td>
</tr>
<tr>
<td>1.5</td>
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</tr>
<tr>
<td>3</td>
<td>-2.3</td>
</tr>
</tbody>
</table>

% difference in CV btw income groups 2 and 1: 22 30 35 40 47 55

% difference in CV btw income groups 3 and 1: 41 55 63 73 85 99

Notes: All columns display results that correspond to the top two panels of the second column of Table 8 (using a 10% import price shock). But instead of setting $\sigma^2_D = \sigma^2_M = 0$, we set $\sigma^2_D = x \times 0.0005$ and $\sigma^2_M = x \times 0.0025$ (as described in footnote 37), for values of $x$ displayed at the top of each column. Column 1 corresponds to column 2 of Table 8.

fare between incomes grow. This growth is entirely driven by the expenditure-switching effect.

**Sensitivity.** We provide a range of additional results in Appendix D. First, we consider import price declines rather than increases. In this case, high-income households benefit more from the first-order effect (they have higher initial import shares) whereas low-income households benefit more from the expenditure-switching effect (they have higher price elasticities). If $\sigma = 0$, the expenditure-switching effect dominates for large import price declines; it also dominates for our smallest import price decline of 2.2% if $\sigma$ is calibrated to match its observed increase in 2014–15. Second, we set $\eta_s = -1.5$, which is at the lower end of our estimates of differences in elasticities of substitution across incomes, rather than $\eta_s = -2$. This slightly reduces differences in welfare across incomes induced by expenditure switching. Third, we vary $\eta_s$ so that the lowest elasticity of substitution (that for the highest-income household, with income of 120,000 CHF) is equal to 1.5 or 5, instead of equal to 3. This leaves the differences in welfare across households largely unchanged (except for the extreme shock to import prices that essentially results in autarky). Fourth, we set the elasticity of substitution between sectors to $\rho = 0.2$ instead of $\rho = 0.99$, which does not have a strong impact on our results. Finally, we consider alternative assumptions on the value of $\eta_s$ in non-grocery sectors $s$. Even when households share common elasticities within the service sector we obtain very similar results.
6 Conclusions

In this paper we revisit a classic question: what are the distributional implications of changes in foreign prices? We focus on differential changes in costs of living across households.

Theoretically, we show that differences across households in compensating variation in response to given income and price changes are shaped by initial expenditure shares across products and initial compensated cross-price elasticities. Empirically, we use detailed Swiss data to document that lower-income households engaged in significantly more expenditure switching towards imported goods in response to the 2015 Swiss franc appreciation. Leveraging these data and imposing generalized, non-homothetic CES preferences, we estimate substantially higher elasticities of substitution for lower-income households.

Import price increases in Switzerland harm higher-income households more than lower-income households both because higher-income households have higher initial import shares (the standard channel considered by the literature) and because they engage in less expenditure switching between imported and domestic goods (a channel from which the literature has abstracted). Quantitatively, we show that for large and dispersed price changes, unequal expenditure switching generates substantial differences in welfare across the income distribution.

Unequal expenditure switching can be relevant for the distributional consequences of high-inflation episodes, if these coincide with a rise in the dispersion of price changes (for evidence on the relation between inflation and price dispersion, see, e.g., Alvarez et al., 2019). We leave this for future research.

References


