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SOURCES OF BIAS IN INFLATION RATES AND IMPLICATIONS FOR INFLATION DYNAMICS

Sarah Lein and Rahel Braun

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JEL Classification: E31, E4, E5, C3, C23

Keywords: Homescan data, Inflation measurement, bias in inflation indexes

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Sources of Bias in Inflation Rates and Implications for Inflation Dynamics*

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January 2021

Abstract

Official statistics measuring the cost of living are known to suffer from several biases. This paper shows that the size of the biases can vary with economic conditions. Using homescan data, it is first confirmed that official price indexes can be tracked using such granular datasets. While the often-acknowledged substitution bias is shown to be relatively small, neglected preference adjustment and product entry/exit results in a 2.6 percentage point bias in the annual inflation rate on average. Furthermore, the bias is particularly large in the aftermath of a shock to relative prices, increasing to 3.7 percentage points.

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1 Introduction

Price indexes are a closely tracked statistics by both the public and policy makers because they measure how the cost of living evolves over time. Growth rates in cost-of-living indexes are therefore an important ingredient in consumers' saving and investment choices and wage negotiations. Letting these indexes grow by a low but positive rate is the key aim of most central banks. It is therefore not surprising that national statistical offices invest substantially in making these statistics as accurate as possible.

The price indexes based on traditional methods of data collection and index definitions that are used in most countries are known to suffer from potential biases because of missing quality adjustments (quality bias) and because of slow and infrequent updates of consumption baskets (substitution bias and product entry/exit bias), as described in Hausman (2003).¹ Furthermore, traditional methods do not take into account that consumers can substitute towards the goods that they prefer. Recently, Redding and Weinstein (2019) show that neglecting these consumer preference adjustments can result in a potentially large bias.

While estimates of the average size of these biases have been documented, some of the biases described above may change in response to economic conditions. In particular, the substitution bias and the consumer valuation bias arise because more traditional price indexes neglect shifts in consumers' expenditures during the periods of observation. We add to this literature by examining how the biases react to a shock to relative prices, where consumers shift consumption baskets more than they do in periods when relative prices do not change systematically. For example, consumers may substitute more intensively towards cheaper or more appealing goods if relative prices within a product category change (even if the average price in the product category remains constant).

In this paper, we quantify the average inflation measurement biases and by how much these biases respond to a shock to relative prices. To do so, we examine the change in relative prices between imports and domestically produced goods observed in Switzerland after the Swiss National Bank abandoned the policy of a minimum exchange rate of CHF 1.20 per EUR in January 2015, as documented in Auer, Burstein, and Lein (2020) and illustrated in

¹See also the Boskin Commission Report (Boskin, Jorgenson, Dulberger, Gordon, and Griliches, 1996).

Figure 1: consumer prices for imports declined by approximately 4% six months after the shock in response to the 14% appreciation of the CHF against the EUR, while domestically produced product's prices remained relatively stable.²

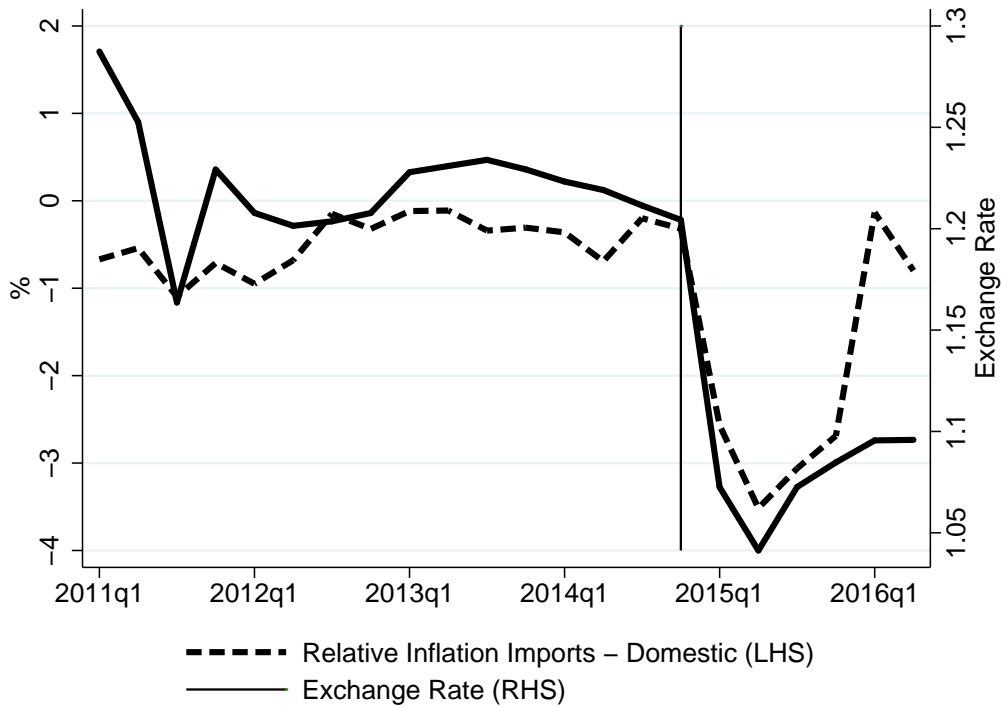
As a result of the decline in relative prices for imports, consumers switch expenditures towards imported products (Auer, Burstein, and Lein, 2020). Biases to official price indexes might be particularly relevant or even larger than in normal times because, in this period, consumers switch expenditures systematically towards products with lower prices. Therefore, the substitution bias may be larger in periods with relative price shifts than during periods when relative prices remain more stable.

Homescan data contain very similar price information as that of the official price index for comparable product categories. We first show that the underlying homescan data compiled by AC Nielsen for Switzerland can replicate official statistics very well if we use exactly the same procedures as those of statistical offices, that is, we keep consumption baskets fixed for a period and let prices vary (Laspeyres-type index).³ We then compare the index that replicates the official statistics to the more flexible Redding-Weinstein CES unified price index (CUPI), which is not subject to the traditional substitution bias, captures consumer preference adjustments, and allows for product entry and exit. By observing the differences between the two indexes, we find largely similar average levels of bias compared to those of the US figures documented in Redding and Weinstein (2019): neglecting consumer valuation results in a bias of 1.3 percentage points (a bit larger than the figures for the US), and neglecting product entry/exit also results in a bias of 1.3 percentage points. Meanwhile, quality bias is already controlled for in homescan data because changes to existing products result in a new product barcode and therefore are included in the variety adjustment bias, we therefore do not touch upon it in our analysis. The traditional substitution bias, which results from using lagged

²The Swiss episode is particularly interesting because it is characterized by a very stable macroeconomic environment before and after the shock; that is, the responses of domestic prices and consumption patterns are arguably largely due to the decline in import prices and not due to other shocks that might affect consumption patterns and relative prices. See Auer, Burstein, and Lein (2020) for a detailed description of this unique episode, and Bonadio, Fischer, and Saure (2018), Kaufmann and Renkin (2017, 2019), Efung, Fahlenbrach, Herpfer, and Krüger (2016), and Auer, Burstein, Erhardt, and Lein (2019), for documentations of the responses of border export and import prices, employment, investment, and export volumes to this exchange-rate shock.

³The only paper we are aware of that calculates inflation rates from homescan data is Kaplan and Schulhofer-Wohl (2017), who calculate inflation rates at the household level and show that the average of these household-level inflation rates varies with the official inflation rate based on the US CPI.

Figure 1



Notes: Relative inflation rates are the difference between inflation of imported products' prices and inflation of domestically-produced products. Inflation rates are calculated with the Jevons formula. See section 3 for more details on index constructions. The exchange rate is the nominal CHF/EUR exchange rate. The vertical line indicates the period before the shock.

weights, is rather small (<0.4 percentage points), also in line with previous literature for the US and other countries.⁴

The size of the consumer valuation and variety adjustment bias increases by an additional 2 percentage points after the shock to relative prices. While the traditional substitution bias remains small in the period of the large shock, consumer valuation and variety adjustment lead to a 1.9 times larger response of the CUPI to the exchange-rate shock than what is suggested by the official price index. The largest contribution of this additional decline comes from the fact that prices in the CUPI are not weighted by expenditure shares but instead are adjusted for preferences: even though consumers are faced with substantially low prices for imported products compared to domestically produced goods, expenditures for domestically

⁴The substitution bias can be reduced by a more frequent collection of weights. Hence, it is estimated to be rather small, that is, approximately 0.4 percentage points per year (Boskin, Jorgenson, Dulberger, Gordon, and Griliches, 1996). For Switzerland, Diewert, Huwiler, and Kohli (2009) estimate the annual substitution bias to be 0.13% on average between the years 1993 and 2002.

produced goods remain high. Therefore, a price index, which weights prices of domestic products by more than prices of imported products, declines by less than the CUPI price index, that does not weight by expenditure shares. There is no larger than usual preference adjustment for domestic and imported goods during the shock period in the CUPI, and imported products account for a larger share in the CUPI than in the expenditure-weighted price index (because imports account for around 42% of the total number of products, but only for 24% of expenditures).

This finding indicates that the bias in inflation rates cannot be corrected for simply by subtracting or adding a constant reflecting the size of the average bias. The dynamics of the biases are also relevant. Therefore, recent attempts to change the measurement of inflation using high-frequency data of online prices, as in Cavallo (2013), or detailed data on prices and purchased quantities, as in Redding and Weinstein (2019), may be a promising approach to accurately measure the levels and dynamics of changes in the cost of living.

This paper is structured as follows. Section 2 outlines the data employed and presents some descriptive statistics. Section 3 computes and compares the different price indexes that are reported in Section 4. Section 5 discusses the results on the backdrop of the literature on optimal inflation rates, and Section 6 concludes.

2 Data and descriptive statistics

The analysis is based on a rich scanner dataset provided by AC Nielsen that contains daily information from over 5,000 households distributed across Switzerland (except Ticino). The database includes information on how much (number and size of packages) and at what price households purchase food, non-alcoholic beverages, alcoholic beverages and tobacco. The dataset covers observations between January 01, 2010 and June 30, 2016. The goods are purchased in stores of 18 different chains, including the top-selling retailers Migros and Coop, as well as gas-station shops, drug stores, outlets and online stores. The different varieties are classified by the European Article Number (EAN) and are divided into supergroups (SGs), product groups (PGs) and product classes (PCs). The following table gives an example of

three observations.

Table 1: *Example of a database entry*

Date	HH	SG	PG	PC	EAN	Shop	P	Q	Size
2011/08/12	1069	dairy	cheese	hard	7610540115229	spec. shop	2.29	1	120 g
2014/02/25	563	bakery	bread	nonperish.	24001148	Aldi	1.25	2	500 g
2015/02/03	3029	afb	juice	nectar	7616835202754	Migros	2.15	1	330 ml

Notes: This table shows an example of a database entry. The prices and EAN codes are randomized, they do not correspond to the true data to comply with data non-disclosure agreements. SG = supergroup, PG = product group, PC = product class, afb = alcohol free beverages. HH= household, P=price per product, Q=quantity of products purchased, Size= package size.

Information on whether a good is imported or domestically produced is added for our further analysis. Therefore, we first search the internet for information on the origin of a good using the classification in Auer, Burstein, and Lein (2020). If this information cannot be found, the EAN helps us to categorize the good. The first two digits of the EAN identify the country in which the manufacturer has registered the product (for Switzerland 76).⁵ We additionally identify goods as imports if it is clear that they do not come from Switzerland even though their EAN begins with 76.⁶ EANs beginning with a 2 indicate an in-store classification. If we cannot determine the origin of in-store goods through the database search, we remove them from the analysis, mainly because stores may re-use them for a different product and because different stores may use the same number for different products, which may affect our analysis.

We remove observations where the package size per EAN is three times larger than the mode or where the price is four times below or above the mean price within this EAN. Furthermore, we consider only households who report more than eight times per year. Some product classes have more than one unit of measurement (e.g., some sauces are measured in grams while others are measured in milliliters). Due to the higher comparability, we keep only observations with the most frequent unit of measurement within a product class. We also drop products that are not included in the official consumer price index (CPI) subgroups

⁵Here we differ from Auer, Burstein, and Lein (2020), as we also include products that are not very clearly identified as domestic or imported through the net search. We do so because we pay less attention to relative prices between domestic and imported products, and we want to keep more products in the sample, which allows us to evaluate product entry and exit and consumer valuation on a broader basis. When we compute the aggregate indexes, the main part of our analysis is not based on a distinction between imports and domestic products.

⁶These products include salmon, seafood, bananas and other exotic fruits, and coffee beans.

food, beverages, and tobacco to make our product sample comparable to the official price index because our first aim is to reconstruct the official index using our data on prices and expenditures.

Finally, we distinguish three samples. First, the *super common varieties sample* $\Omega_{\forall t}$. Here, we keep only goods that are purchased at least once every quarter. This is most in line with the collection procedure of the SFSO (Swiss Federal Statistical Office) since they also observe a fixed amount of goods, and new products are usually entered with a lag.⁷ Second, the *common varieties sample* $\Omega_{t-1,t}$, where we include all goods that have been observed in a quarter t and the same quarter in the previous year. Third, the *full sample* Ω_t includes all goods, independent of the entry and exit date. Table 2 shows the transactions and number of products in each dataset.

We harmonize the classification provided in the raw data from AC Nielsen such that the product categories are better comparable to the SFSO classification. From henceforward, whenever we talk about a *productclass* n we mean the lowest index position, while a *variety* k means the unique product (EAN). For example, for the *productclass* rice, and the *varieties* are different brands or types of rice, such as "Uncle Ben's Gold Premium Basmati" or "M-Classic Carolina Dried Rice".

Table 3 gives an overview over the different expenditure categories. The first column of the table, the *SFSO share*, reports the mean weight in the calculations of the CPI based on the official annual personal consumption expenditure survey for the years 2010-2015. The second column, labeled *share*, is the equivalent expenditure share in our data, which is calculated as the expenditure for each group over total expenditure. The two columns are largely similar (the correlation is 0.7), which suggests that scanner data capture the same information as the data collected by the SFSO. The third column reports the *Import ratio*, which is the expenditure for imports over total expenditure within a group.

The four categories we cover in the scanner dataset, namely, food, non-alcoholic and alcoholic beverages, and tobacco, account for roughly 13% of the total official consumer price

⁷The SFSO collects prices at a monthly frequency; however, the results do not change significantly. See Swiss Federal Statistical Office (2016) for details about the SFSO calculations

Table 2: *Data overview*

	Transactions	Expenditure Share (%)	Products
<i>Full Sample: Ω_t</i>			
Total	13,854,904	100	116,766
Swiss	10,682,048	75.45	59,447
Imports	3,172,856	24.55	57,319
Imports EA	2,384,804	17.52	49,129
Imports ROW	788,055	7.03	8,191
<i>Common Varieties: $\Omega_{t-1,t}$</i>			
Total of Ω_t	10,284,137	74.23	60,934
Swiss	7,945,735	75.86	35,674
Imports	2,338,402	24.14	25,260
Imports EA	1,743,861	17.09	21,259
Imports ROW	594,542	7.06	4,002
<i>Super Common Varieties: $\Omega_{\forall t}$</i>			
Total of Ω_t	7,366,730	53.18	6,695
Swiss	5,782,141	76.76	4,687
Imports	1,584,589	23.24	2,008
Imports EA	1,166,444	16.34	1,506
Imports ROW	420,045	6.90	502

Notes: Transactions are the number of purchases observed, Expenditure Share is the share of total expenditures in the respective sample (in %), and Products shows the number of unique products. Swiss goods are produced and sold in Switzerland, imports are sold but not produced in Switzerland, imports EA denote imports from the euro area and imports ROW are imports from outside the euro area.

The full sample includes all varieties, the common sample only varieties sold in period t and period $t - 1$, and the super common varieties sample includes only products that are sold in all periods observed.

index. ⁸

⁸See Swiss Federal Statistical Office (2016).

Table 3: *Product categories in CPI and homescan data*

	<i>SFSO share</i>	<i>share</i>	<i>import ratio</i>
	%	%	%
Beer	1.20	1.47	22.45
Bread, Flour	13.15	12.32	13.43
Cigarettes	4.95	5.97	11.31
Coffee, Tea	3.25	3.68	14.01
Cooking Fat, Oil	2.11	3.33	15.35
Fish	2.83	0.99	21.26
Fruits and Vegetables	15.50	10.07	17.00
Meat	19.59	8.78	7.53
Milk, Cheese	13.40	18.40	16.04
Other Aliments	5.76	11.16	20.60
Other Tobacco	0.30	0.19	4.45
Softdrinks	4.55	6.60	31.18
Spirits	1.17	0.98	65.28
Sugar, Sweets	5.44	12.11	22.92
Wine	6.82	3.91	47.37
Total	6.67	6.66	22.01

Notes: This table shows the weights of the different groups for the official data from the SFSO (second row) and the scanner dataset (third row). The fourth row shows the import ratio in the respective group in the scanner dataset.

3 Price Indexes

In this section, we first summarize the traditional Laspeyres price index. It is based on lagged consumption expenditures, which are used to weight the current period prices. Since consumers tend to substitute more expensive for cheaper goods, the Laspeyres index overestimate inflation, leading to a positive substitution bias. Another widely used concept to measure changes in prices is the Paasche index, which uses end-of-period weights and therefore underestimates inflation. The Fisher index, which is a geometric average of the Laspeyres and Paasche indexes, is usually referred to as an ‘exact index’ because it minimizes substitution bias. However, the Fisher index does not allow for preference adjustments and does not consider the utility gain from net product entry. The CUPI developed in Redding and Weinstein (2019) overcomes these limitations: it is much more flexible and nests the indexes discussed above.

3.1 Laspeyres, Paasche, and Fisher Price Indexes

The formula for the Laspeyres (1871) index is

$$\Phi_t^{Las} = \left[\sum_{k \in \Omega_{t,t-1}} S_{kt-1} \left(\frac{p_{kt}}{p_{kt-1}} \right) \right]$$

with $S_{kt-1} = \frac{p_{kt-1}q_{kt-1}}{\sum_l p_{lt-1}q_{lt-1}}$,

where $t - 1$ is any period in the past. p_{kt}/q_{kt} is price/quantity of variety k purchased at time t .

The Paasche (1874) index uses weights from the current period t instead of a previous period $t - 1$.

$$\Phi_t^{Paa} = \left[\sum_{k \in \Omega_{t,t-1}} S_{kt} \left(\frac{p_{kt}}{p_{kt-1}} \right)^{-1} \right]^{-1}$$

with $S_{kt} = \frac{p_{kt}q_{kt}}{\sum_l p_{lt}q_{lt}}$.

Because the weights already include substitution, this index tends to underestimate inflation.

The Fisher index, calculated as the geometric average of the Paasche and Laspeyres indexes, suffers only minimally from substitution bias

$$\Phi_t^{Fis} = (\Phi_t^{Paa} \cdot \Phi_t^{Las})^{0.5}.$$

3.2 CES Unified Price Index

Recently, Redding and Weinstein (2019) proposed a new method to calculate price indexes that takes into account both changes in the demand for individual goods and the entry and exit of goods over time. They show that both aspects play an important role and that including them lowers the cost of living substantially.

The CES Unified Price Index (CUPI) is based on the economic approach of inflation calculation, which assumes a specific utility function for households, and inflation is the

change in the cost of living while keeping utility constant.⁹ Suppose that households have constant elasticity of substitution (CES) preferences in the following form:

$$\mathbb{U}_t = \left[\sum_{k \in \Omega_{t,t-1}} (q_{kt})^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}}$$

where σ is the elasticity of substitution. Household optimization results in the Sato-Vartia index given by:

$$\Phi_{t-1,t}^{SV} = \prod_{k \in \Omega_{t,t-1}} \left(\frac{p_{kt}}{p_{kt-1}} \right)^{\omega_{kt}^*}$$

with $\omega_{kt}^* \equiv \frac{\frac{S_{kt}^* - S_{kt-1}^*}{\ln(S_{kt}^*) - \ln(S_{kt-1}^*)}}{\sum_{l \in \Omega_{t,t-1}} \frac{S_{lt}^* - S_{lt-1}^*}{\ln(S_{lt}^*) - \ln(S_{lt-1}^*)}}$

Redding and Weinstein (2019) expand this approach with two features. They also assume constant elasticity of substitution (CES) preferences¹⁰

$$\mathbb{U}_t = \left[\sum_{k \in \Omega_t} (\varphi_{kt} q_{kt})^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}}$$

however, they add preference (taste) parameters $\varphi_{kt} > 0$ and include new and exiting products Ω_t rather than only varieties bought in both periods $\Omega_{t,t-1}$ ¹¹. Household optimization then

⁹The inflation calculation with the economic approach and the resulting cost-of-living index goes back to Konüs (1924). Ever since, many other authors have contributed to this approach, including Diewert (1976), who shows that the Fisher index approximates the cost-of-living index resulting from the economic approach, Neary (2004), who discusses how to use the economic approach to compare price levels across countries, and Feenstra (1994), who shows how inflation rates calculated with the economic approach can incorporate new and exiting goods. Indexes such as the Lloyd (1975)-Moulton (1975) index and the Sato (1976)-Vartia (1976) index are based on the economic approach.

¹⁰The authors show that the main results also hold for other preference functions.

¹¹The authors base these results on Feenstra (1994), who originally demonstrated how to incorporate new and exiting products.

yields the following unit expenditure function ¹²

$$\mathbb{P}_t = \left[\sum_{k \in \Omega_t} \left(\frac{p_{kt}}{\varphi_{kt}} \right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}}$$

and expenditure shares (applying Shepard's Lemma)

$$S_{kt} \equiv \frac{p_{kt}q_{kt}}{\sum_{l \in \Omega_t} p_{lt}q_{lt}} = \frac{(p_{kt}/\varphi_{kt})^{1-\sigma}}{\sum_{l \in \Omega_t} (p_{lt}/\varphi_{lt})^{1-\sigma}}. \quad (1)$$

The cost of living is defined as the change in the unit expenditure function from period t to $t - 1$.

$$\Phi_{t-1,t} = \frac{\mathbb{P}_t}{\mathbb{P}_{t-1}} = \left[\frac{\sum_{k \in \Omega_t} (p_{kt}/\varphi_{kt})^{1-\sigma}}{\sum_{k \in \Omega_{t-1}} (p_{kt-1}/\varphi_{kt-1})^{1-\sigma}} \right]^{\frac{1}{1-\sigma}} \quad (2)$$

To account for the role of product entry and exit, Redding and Weinstein (2019) decompose equation (2) into a unit expenditure function on common varieties only ($\Omega_{t,t-1}$) and a term that captures the variety of new entries and exits between t and $t - 1$ (variety adjustment term). λ_t expresses expenditures on all goods available in both periods (common varieties) as a share of total expenditures in t

$$\lambda_t \equiv \frac{\sum_{k \in \Omega_{t-1,t}} (p_{kt}/\varphi_{kt})^{1-\sigma}}{\sum_{k \in \Omega_t} (p_{kt}/\varphi_{kt})^{1-\sigma}},$$

which is the total expenditure of continuing varieties divided by the expenditure on all varieties available in t evaluated at current prices. This value decreases if the share of new varieties increases.

¹²Section B of the online appendix shows the derivation

Similarly,

$$\lambda_{t-1} \equiv \frac{\sum_{k \in \Omega_{t-1,t}} (p_{kt-1}/\varphi_{kt-1})^{1-\sigma}}{\sum_{k \in \Omega_{t-1}} (p_{kt-1}/\varphi_{kt-1})^{1-\sigma}},$$

is the total expenditure share of continuing varieties (as a share of past total expenditures) in $t-1$ prices. This value decreases if the share of exiting varieties increases.

Hence, equation (2) can be written as

$$\begin{aligned} \Phi_{t-1,t} &= \frac{\mathbb{P}_t}{\mathbb{P}_{t-1}} \\ &= \left[\frac{\lambda_{t-1} \sum_{k \in \Omega_{t-1,t}} \left(\frac{p_{kt}}{\varphi_{kt}} \right)^{1-\sigma}}{\lambda_t \sum_{k \in \Omega_{t-1,t}} \left(\frac{p_{kt-1}}{\varphi_{kt-1}} \right)^{1-\sigma}} \right]^{\frac{1}{1-\sigma}} \\ &= \left(\frac{\lambda_{t-1}}{\lambda_t} \right)^{\frac{1}{1-\sigma}} \frac{\mathbb{P}_t^*}{\mathbb{P}_{t-1}^*} \end{aligned} \quad (3)$$

where $\mathbb{P}_t^* \equiv \left[\sum_{k \in \Omega_{t-1,t}} \left(\frac{p_{kt}}{\varphi_{kt}} \right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}}$ is the unit expenditure function defined over the set of common varieties $\Omega_{t-1,t}$, and $\frac{\mathbb{P}_t^*}{\mathbb{P}_{t-1}^*}$ is the CES exact price index for common varieties (CVV).

To make the price index money-metric, Redding and Weinstein (2019) assume a constant geometric mean of all demand parameters $\bar{\varphi} \Rightarrow \frac{1}{N_{t,t-1}} \sum_{k \in \Omega_{t,t-1}} (\ln(\varphi_{kt}) - \ln(\varphi_{kt-1})) = 0$.

Rewriting equation (3) yields the CUPI ¹³

$$\Phi_{t-1,t}^{CUPI} = \underbrace{\left(\frac{\lambda_t}{\lambda_{t-1}} \right)^{\frac{1}{\sigma-1}}}_{\text{Variety Adjustment}} \left[\underbrace{\frac{\tilde{P}_t^*}{\tilde{P}_{t-1}^*}}_{\text{Jevons}} \underbrace{\left(\frac{\tilde{S}_t^*}{\tilde{S}_{t-1}^*} \right)^{\frac{1}{\sigma-1}}}_{\text{Preference Adjustment}} \right] \quad (4)$$

where a tilde represents the geometric average and an asterisk denotes the set of common varieties such that $\tilde{x}^* = \left(\prod_{k \in \Omega_{t,t-1}} x_{kt} \right)^{\frac{1}{N_{t,t-1}}}$. $N_{t,t-1}$ is the quantity of common varieties available in t and $t-1$.

¹³For the detailed derivation see Redding and Weinstein (2019) and their web appendix.

Equation (4) shows that the CUPI is the product of three terms: the variety adjustment term, the ratio of two Jevons price indexes for common varieties at time t and $t - 1$, and the preference adjustment for common varieties. The first term measures how much consumers value new goods relative to exiting goods. If new goods have lower prices (adjusted for the preference parameter) than those of the exiting goods, the price index falls because the new variety is more appealing to consumers and they therefore obtain greater utility for a given amount of expenditures. The second term is simply an unweighted geometric average (Jevons index) of all prices that are observable in $t - 1$ and t . The last term is a function of the geometric average of expenditure shares on common varieties, \tilde{S}_t^* . If this term decreases, that is $\tilde{S}_t^* < \tilde{S}_{t-1}^*$, expenditure shares become more uneven across varieties. The elasticity of substitution is larger than one; hence, consumers value that they can buy larger quantities of varieties they find more appealing and small quantities of others. Thus, a larger heterogeneity in these expenditure shares is associated with a higher utility and therefore with a smaller unit expenditure function.¹⁴

The estimation of the elasticity of substitution is based on the methodology by Feenstra (1994), Broda and Weinstein (2006), and Broda and Weinstein (2010), and outlined in section C of the online appendix. Our estimates of the elasticity of substitution yields $\sigma = 3.9$, close to estimates of demand elasticity reported in the literature using similar micro data.¹⁵

4 Results

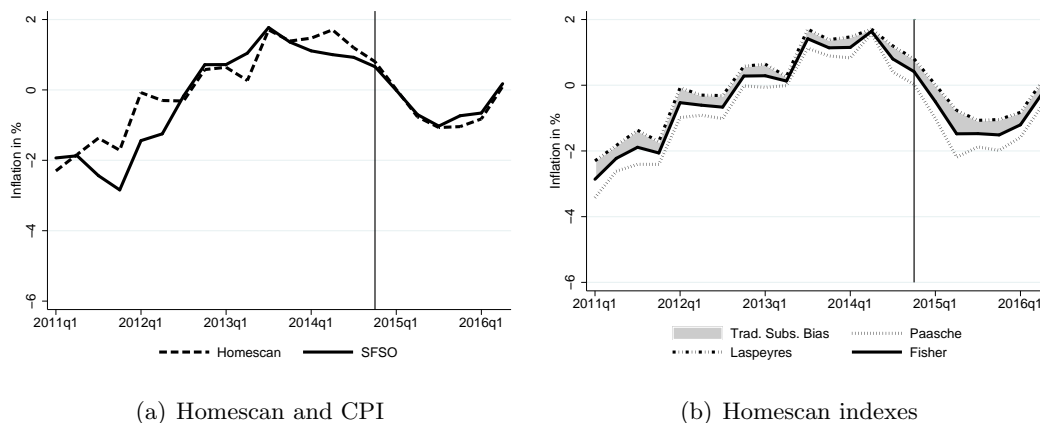
In this section, we first show that homescan data can be used to replicate the official CPI calculated by the SFSO. We then show how much the Sato-Vartia price index deviates from the CUPI, where the largest contribution off these differences comes from, and whether the contributions change after the price shock.

¹⁴This expenditure shifts term is only negligible if either the elasticity of substitution is infinity, that is, goods are perfect substitutes, or if all preference-adjusted prices are identical and therefore all expenditure shares are equal across varieties. When preference-adjusted prices become more dispersed, so do expenditure shares across varieties. Because the expenditure shares \tilde{S}_t^* are geometric averages, this results in a decline in the cost-of-living index, see Redding and Weinstein (2018).

¹⁵See for example, Redding and Weinstein (2019), Beck and Lein (2020), or DellaVigna and Gentzkow (2019)

4.1 Comparison with the official index

Figure 2



The SFSO series is the CPI for the subcomponents food, beverages, and tobacco. The official weights for each of these subcomponents are re-adjusted such that they sum to one. Sources: SFSO, own calculations.

Panel (a) of Figure 2 shows the CPI calculated with the homescan dataset and the official CPI for food, beverages and tobacco.¹⁶ The official CPI figure (we always compare with the index for the same product categories not the overall index including services, etc.) is very close to our homescan-based index. This suggests that the information included is very similar, which is important for the subsequent analysis, when we allow for a more flexible definition of the consumer price index.

4.2 Bias decomposition

Substitution bias. Having shown that our scanner dataset contains the same information as that which the SFSO uses for the calculation of the official CPI, we examine whether the official CPI is biased due to a substitution effect caused by the lagged weights. Therefore, we exploit the simultaneity of the scanner dataset (i.e., we observe weights and prices at the same time) and calculate the Laspeyres index with lagged weights, the Paasche index with contemporaneous weights, and the Fisher index, which minimizes the substitution bias. The difference between the Laspeyres and the Fisher index is typically used as a measure of the

¹⁶We used the the Laspeyres index to calculate the homescan inflation rate. However, the SFSO actually uses the Young index which is close to the Laspeyres index and is outlined in Section A of the online appendix

traditional substitution bias. To calculate the difference, we focus on the common varieties sample ($\Omega_{t-1,t}$), that is, we do not allow for entry and exit within two time periods.

The influence of lagged weights is shown in Panel (b) of Figure 2. The average substitution bias is approximately 0.38 percentage points and is relatively stable over time (gray area). In 2015, the bias increases to 0.52 percentage points, which is still relatively small. Hence, even after the large exchange rate shock, the inflation rates are only slightly lower when considering the contemporaneous weights. The bias found here is in line with that in previous studies for the US (Boskin, Jorgenson, Dulberger, Gordon, and Griliches (1996), Kaplan and Schulhofer-Wohl (2017)) and slightly higher than the one for Switzerland (Diewert, Huwiler, and Kohli, 2009). The small substitution bias is also in line with Hausman (2003), who shows that this is a second-order bias. Intuitively, the bias is the change in expenditures multiplied by the change in prices, and these values are both small numbers, even after a large shock.

Having shown that the traditional substitution bias is small, we examine the impact of using expenditure-weighted prices, instead of prices adjusted for preference parameters, and of using a common varieties sample, instead of a sample that allows for product entry and exit. To quantify the bias in inflation levels and the contributions of each of the sources of bias, we decompose the CUPI into its three terms.

$$\ln(\Phi_{t-1,t}^{CUPPI}) = \underbrace{\frac{1}{\sigma-1} * \ln\left(\frac{\lambda_t}{\lambda_{t-1}}\right)}_{\text{Variety Adjustment}} + \underbrace{\ln\left(\frac{\tilde{P}_t^*}{\tilde{P}_{t-1}^*}\right)}_{\text{Jevons}} + \underbrace{\frac{1}{\sigma-1} * \ln\left(\frac{\tilde{S}_t^*}{\tilde{S}_{t-1}^*}\right)}_{\text{Preference Adjustment}} \quad (5)$$

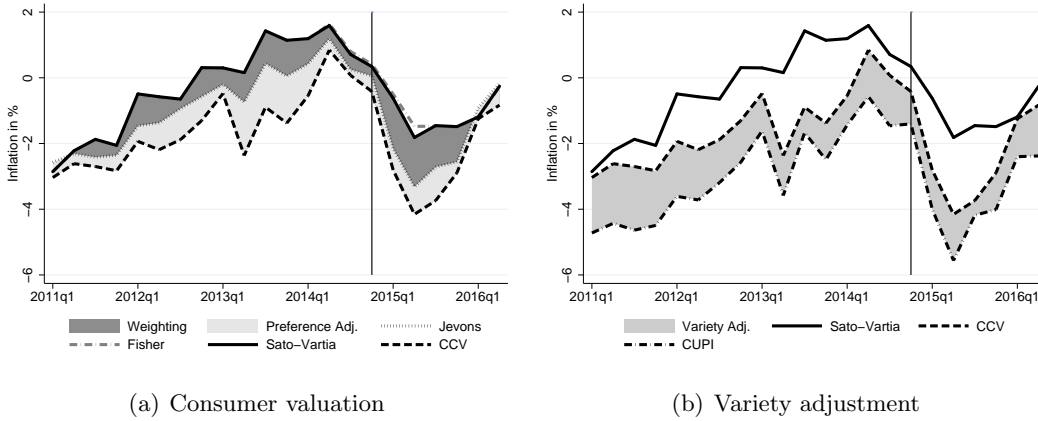
Subtracting the Sato-Vartia index on both sides of the equation yields the bias

decomposition

$$\begin{aligned}
 \ln(\Phi_{t-1,t}^{CUI}) - \ln(\Phi_{t-1,t}^{SV}) &= \underbrace{\frac{1}{\sigma-1} * \ln\left(\frac{\tilde{S}_t^*}{\tilde{S}_{t-1}^*}\right)}_{\text{Preference Adjustment}} + \underbrace{\sum_{k \in \Omega_{t,t-1}} \left(\frac{1}{N_{t,t-1}} - \omega_{kt}^*\right) \ln\left(\frac{p_{kt}}{p_{kt-1}}\right)}_{\text{Weighting}} \\
 &\quad \underbrace{\hspace{10em}}_{\text{Consumer Valuation}} \\
 &+ \underbrace{\frac{1}{\sigma-1} * \ln\left(\frac{\lambda_t}{\lambda_{t-1}}\right)}_{\text{Variety Adjustment}} \tag{6}
 \end{aligned}$$

which is the sum of two terms. The first term is the consumer valuation bias, the second term the variety adjustment bias. The first term can be further decomposed into the term preference adjustment (explained above in section 3) and the term that we label ‘weighting’, because it quantifies the role of using a price index weighted by expenditure shares (the Sato-Varita index) and the unweighted index (Jevons).

Figure 3: *The role of weighting, expenditure dispersion, and variety adjustment*



This figure shows the weighting and preference adjustment bias (Panel (a)) as well as the variety adjustment bias (Panel (b)) together with Jevons, Fisher, Sato-Vartia, CCV, and CUI.

Weighting bias. Panel (a) of Figure 3 plots the Sato-Vartia index as a solid line, the Jevons index as a dotted line, and the CCV as a dashed line.¹⁷ The bias shown as dark grey area comes from the difference between the Sato-Vartia index and Jevons index and is labeled ‘weighting’ because the Sato-Vartia uses expenditure-adjusted weights while the Jevons uses equal weights for different varieties. This difference is 0.6 percentage points on average.

¹⁷Note that the Sato-Vartia index and the Fisher index are almost the same.

Preference adjustment bias. The difference between the Jevons index and the CCV, is shown as light grey area in Panel (a) of Figure 3. It captures how much dispersion of expenditures across products change and therefore the heterogeneity in prices *adjusted for preferences*. Also this component of the total bias shifts the inflation rate downwards, since expenditures become more and more dispersed over time (see section 3). This bias amounts to 0.7 percentage points on average. Thus, taking the weighting and preference adjustment together, the total consumer valuation bias from using expenditure weighted instead of preference adjusted prices amounts to 1.3 percentage points in total.

Variety adjustment bias. Finally, we quantify the role of net product entry by including the variety adjustment term from equation (4). The bias arising from neglecting this term is substantial. Panel (b) of Figure 3 shows that controlling for variety adjustment inflation down by 1.3 percentage points, which is a large and persistent bias.

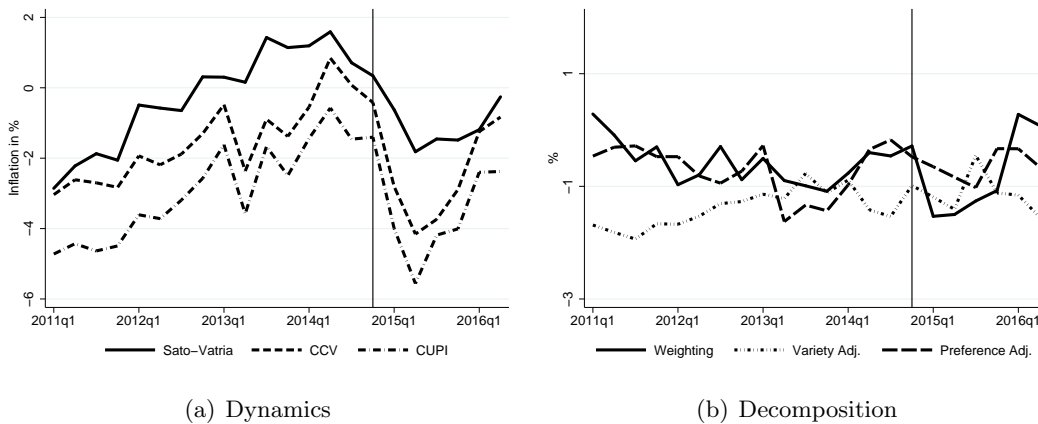
4.3 The role of biases in inflation dynamics

We have shown that, on average, neglecting consumer valuation results in a bias of 1.3 percentage points and neglecting product entry/exit results in another 1.3 percentage point bias. In this section, we study whether the biases differ after relative price changes. Therefore, in Panel (a) of Figure 4, we plot the main indexes again and focus on the time after the exchange rate shock in January 2015. Panel (b) shows the three components of the bias separately (as in equation (6)).

The index which includes the preference shocks and variety adjustment (CUPI) clearly declines much more than the index not including them (Sato-Vartia). The CUPI declines by 4.2 percentage points between 2014Q4 and 2015Q2, while the Sato-Vartia index declines by only 2.2 percentage points. In the following, we study why the CUPI reacts more to the shock.

The preference adjustment term and the variety adjustment term do not react to the shock. However, the Jevons index (and hence the weighting term) reacts strongly (see Panel (b)). As the Jevons index uses uniform weights, this suggests that varieties with larger price declines tend to be those with relatively smaller expenditure weights in the Sato-Varita index. In this

Figure 4



This figure shows the Laspeyres, Jevons, Fisher, Sato-Vartia, CCV and CUPI, as well as the Variety Adjustment and Expenditure Shifts Bias defined below in equation (6).

specific case, after the exchange rate shock, the prices of Swiss-produced goods decline by much less than those of import goods. In the common varieties sample, 42% of the products are imports, but they account for only approximately 24% of expenditures (Table 2). This explains why the Sato-Varita decline by less: import prices decline by more but account for only 24%, compared to 42% for the Jevons index.

In the CUPI, products are not weighted by expenditure shares, as in the Sato-Vartia index, because a larger weight for a product has a different interpretation. Consider an example where we have only two products and one consumer and the price of one product falls relative to the price of the other product, which has a constant price. If the consumer does not adjust quantities purchased based on the decline in the relative price, the expenditure share of the product with a constant price goes up. The *price adjusted for preference* of this product declines because a higher expenditure share implies that a consumer increased their preference for this product; therefore, the preference-adjusted price goes down even though the unadjusted price remains constant. In this case, consumers continue to purchase domestically produced goods even though their relative price increases. The large weight on domestically produced goods (approximately 76% in our data and the official CPI), whose prices hardly change in response to the shock, mutes the response of the official Sato-Varita price index to the exchange-rate appreciation. In the CUPI, however, the fact that expenditures on

products with rising relative prices do not shift substantially suggest that consumers have stronger preferences for these products and, therefore, that prices for domestically produced goods, adjusted for preferences, decline by much more than suggested by the unadjusted prices.

The main results are summarized in Table (4). Before the shock, the Sato-Vartia and the CUPI differ by 2.6 percentage points on average (the standard deviation over time is 0.6). Almost half of the difference results from consumer valuation, and the other half (52%) results from variety adjustment. However, this changes after the shock. The difference between the Sato-Vartia and the CUPI increases significantly, by 2.0 percentage points, and 61% of this difference comes from the difference between expenditure-weighted prices (which decline by less because of the larger weight on domestic goods, see above), while only 18% come from preference adjustments, and 21% come from variety adjustment.

Table 4: *Bias decomposition*

	2011-2014	Δ 2014Q4-2015Q2
Sato-Vartia-CUPI	2.63 (0.61)	2.00
shares		
Consumer Valuation	0.48	0.79
Weighting	0.21	0.61
Preference Adjustment	0.26	0.18
Variety Adjustment	0.52	0.21

Notes: The first line shows the percentage point difference between the Sato-Vartia index and the CUPI, on average, for 2011-2014 (first column) and between 2014Q4 and 2015Q2 (second column). The shares are calculated as $1 = \frac{Weighting}{Sato-Vartia-CUPI} + \frac{PreferenceAdjustment}{Sato-Vartia-CUPI} + \frac{VarietyAdjustment}{Sato-Vartia-CUPI}$. The number in parentheses in the line Sato-Vartia-CUPI is the standard deviation of the average bias.

5 Implications for the optimal rate of inflation

According to our results, and in line with the findings in Redding and Weinstein (2019), neglecting preference adjustment and the entry/exit of goods results in a 2.6 percentage point lower inflation rate at an annual basis. Thus, official inflation rates, which do not incorporate these two biases, would overstate the true rate of inflation. This might imply that policy makers should target a positive inflation rate. In this section, we briefly discuss our findings

on the backdrop of the literature on optimal inflation rates and show that only the variety adjustment bias justifies a positive inflation target.

The role of consumer valuation bias in optimal inflation rates is closely related to the role of the quality bias. It has long been argued that one motivation for setting positive inflation targets is that the consumer price index overstates the true rate of inflation. Schmitt-Grohé and Uribe (2011) convincingly show that the optimal inflation rate might still be zero, even in the presence of unmeasured quality improvements. We show in section D of the online appendix, that the preference parameters in the derivation of the CUPI can also be interpreted as the quality adjustment term in Schmitt-Grohé and Uribe (2011). In other words, a change in preferences can be interpreted as an unobserved quality change of a good. Therefore, the role of quality adjustment for optimal inflation rates also applies to consumer valuation adjustment. Schmitt-Grohé and Uribe (2011) show that whether the inflation target should be adjusted depends critically on what prices are assumed to be sticky. If prices are set in non-quality-adjusted terms, and thus these prices are sticky, then inflation targets should not be corrected and the optimal inflation rate remains at zero because inefficiencies resulting from price distortions are mitigated. On the other hand, when quality-adjusted prices are sticky, the optimal (non-quality-adjusted) inflation rate should be corrected for the bias since it is efficient to keep quality-adjusted prices constant while non-quality-adjusted prices are allowed to fluctuate.

In the model underlying the CUPI, where the quality adjustment term can be interpreted as consumer valuation, it is reasonable to assume firms' prices that are sticky rather than firms' prices adjusted for preferences. Therefore, the downward bias in inflation due to consumer valuation would not justify an even higher inflation target if central banks observe a Fisher- or Laspeyres-type price index.

Bilbiie, Fujiwara, and Ghironi (2014) show that in a model including endogenous firm entry and price adjustment costs, the optimal inflation rate still equals zero if utility functions have constant elasticity of substitution preferences, as in Dixit-Stiglitz.¹⁸ If statistical offices do not

¹⁸In the case of non-CES preferences, this conclusion is no longer valid because the benefit of variety for society and the incentive for product creation (net markup) do not coincide. Optimal long-run inflation should be positive if there are too many products on the market (Bilbiie, Fujiwara, and Ghironi, 2014).

account for product entry and exit, the optimal inflation rate a central bank might want to target should be adjusted for entry/exit bias. Thus, it is reasonable to set a positive inflation target equal to the size of the variety adjustment bias.

6 Conclusion

In this paper, we evaluate and quantify the extent of bias in the price indexes that are typically used by official statistical agencies to measure the cost of living. A typical index is based on a monthly collection of prices for unique products, and these prices are weighted by expenditure shares from a past period. Thus, the price index documents the evolution of the price level for a fixed basket of goods over time. Price indexes are prone to biases, which include the traditional substitution bias arising from consumers substituting towards cheaper products, the consumer valuation bias arising from neglecting preference adjustment and using expenditure-weighted averages of prices, and the variety adjustment bias arising from ignoring product entry and exits.

We show that all biases are negative, that is, the official inflation rates overestimate true inflation. The traditional substitution bias is relatively small, in the range of 0.4 percentage points per year, whereas the consumer valuation bias and the variety adjustment bias are each approximately 1.3 percentage points per year. These biases are substantial and similar to estimates reported in Redding and Weinstein (2019) for the US.

We add to this literature by examining whether these biases increase after a large shock to relative prices. To do so, we use AC Nielsen homescan data for Switzerland around the 2015 appreciation. The shift in the exchange rate led to a decline in the prices of imported products, while the prices of domestically produced products decreased only marginally. This scenario gives consumers an incentive to substitute and can also change the dynamics for product entry if domestic and foreign firms change the supply in the market due to the increase in competitiveness of foreign products. Furthermore, the change in relative prices and product entry may also shift expenditures of consumers and therefore change the consumer valuation bias.

We find that the traditional substitution bias increases only slightly after the exchange rate shock. More importantly, the CUPI-inflation rate, which takes into account all biases, declines by 4.2 percentage points six months after the exchange rate shock, while the Sato-Vartia inflation rate declines by only 2.2 percentage points, suggesting that there is not just a level shift in these biases. The biases also affect inflation dynamics, as after the shock, the CUPI declines 1.9 times more than does the Sato-Vartia index. These biases also translate into the calculation of real figures. For example, the growth in real private consumption in 2015 was 1.8% in Switzerland (compared to 1.2% in 2014). Adjusting these figures for the difference between the Sato-Vartia index and the CUPI using a simple back-of-the-envelope calculation would result in a much larger increase in real private consumption after the appreciation (4.9% in 2015 and 3.4% in 2014).¹⁹

¹⁹Because the consumption deflator includes more product categories than those in our Sato-Vartia and CUPI indexes, we calculate the growth rate of the consumption deflator and subtract the differences between the Sato-Vartia index and CUPI reported in the paper (adjusted deflator). We then calculate the real consumption growth rates reported above as the difference between the growth rate of nominal consumption and the adjusted deflator.

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Online Appendix

A Young Index

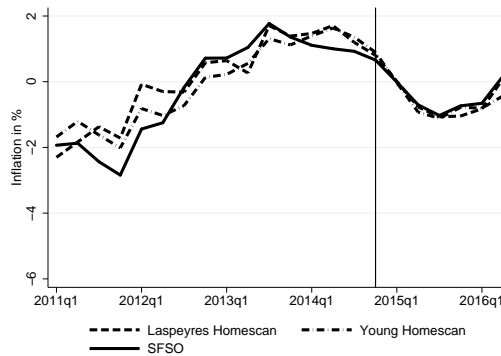
This part of the online appendix shows how the SFSO computes the CPI. The method is similar to the Laspeyres index and is called the Young (1812) index. The formula is:

$$\Phi_t^{You} = \left[\sum_{j \in \Omega_{t,t-1}} S_{jb} \left(\frac{p_{jt}}{p_{jt-1}} \right) \right]$$

with ²⁰ $S_{jb} = \frac{p_{jb}q_{jb}}{\sum_l p_{lb}q_{lb}}$ (7)

The SFSO collects the weights in an annual expenditure survey,²¹. Since the evaluation of this survey takes time, the weights lag by two years. For example, for the CPI of 2015, the weights are from 2013. In the Young Index, this translates to $b = t - 8$ ²² in equation (7). The SFSO aggregates the individual product prices to predefined productclasses n by taking the geometric average. Price changes are calculated relative to 2010q1; hence, p_{nt-1} is the price of productclass n in period 2010q1. The Young Index is very close to the Laspeyres Index as the following figure shows.

Figure 5



The SFSO series is the CPI for the subcomponents food, beverages, and tobacco. The official weights for each of these subcomponents are re-adjusted such that they sum to one. Sources: SFSO, own calculations.

²¹See Swiss Federal Statistical Office (2013)

²²Since the scanner dataset starts in 2010, we cannot calculate weights for the years 2008 and 2009. Therefore, in 2010 and 2011, the lagged weights are from 2010.

B Derivation of the CES Unified Price Index

$$U_t = \left[\sum_{k \in \Omega_t} (\varphi_{kt} q_{kt})^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}}$$

Budget constraint

$$\sum_{k \in \Omega_t} p_{kt} q_{kt} = X_t$$

The cost minimizing problem is:

$$\begin{aligned} \min_{q_{kt}} \quad & \sum_{k \in \Omega_t} p_{kt} q_{kt} \\ \text{st} \quad & \left[\sum_{k \in \Omega_t} (\varphi_{kt} q_{kt})^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}} = U_t^0 \end{aligned}$$

First order conditions:

$$\begin{aligned} \frac{p_{kt}}{p_{it}} &= \left(\frac{q_{kt} \varphi_{kt}}{q_{it} \varphi_{it}} \right)^{\frac{-1}{\sigma}} \frac{\varphi_{kt}}{\varphi_{it}} \\ q_{kt} &= q_{it} \frac{\varphi_{it}}{\varphi_{kt}} \left(\frac{p_{kt}/\varphi_{kt}}{p_{it}/\varphi_{it}} \right)^{-\sigma} \end{aligned}$$

Plugging this back into the utility function gives and solving for q_{it} yields:

$$q_{it} = U_t^0 \frac{1}{\varphi_{it}} (p_{it}/\varphi_{it})^{-\sigma} \left[\sum_{k \in \Omega_t} (p_{kt}/\varphi_{kt})^{1-\sigma} \right]^{\frac{\sigma}{1-\sigma}}$$

which is the Hicksian demand. To obtain the expenditure function, plug the Hicksian demand into the expenditures.

$$\begin{aligned}
\sum_{i \in \Omega_t} p_{it} q_{it} &= \sum_i p_{it} \left(U^0 \frac{1}{\varphi_{it}} (p_{it}/\varphi_{it})^{-\sigma} \left[\sum_{k \in \Omega_t} (p_{kt}/\varphi_{kt})^{1-\sigma} \right]^{\frac{\sigma}{1-\sigma}} \right) \\
&= U^0 \left[\sum_{k \in \Omega_t} (p_{kt}/\varphi_{kt})^{1-\sigma} \right]^{\frac{1}{1-\sigma}}
\end{aligned}$$

Hence, the expenditure function is

$$U^0 \left[\sum_{k \in \Omega_t} \left(\frac{p_{kt}}{\varphi_{kt}} \right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}}$$

C Estimation of the elasticity of substitution

This section explains briefly how the elasticity of substitution is calculated. For more details see Feenstra (1994), Broda and Weinstein (2006), Broda and Weinstein (2010), and Redding and Weinstein (2019).

Suppose that there is the following demand equation

$$\Delta \ln s_{kt} = \psi_t^1 - (\sigma - 1) \Delta \ln(p_{kt}) + \varepsilon_{kt}$$

The error term ε_{kt} includes the preference shocks. The supply equation is:

$$\Delta \ln s_{kt} = \psi_t^2 + \frac{\omega}{1 + \omega} \Delta \ln(p_{kt}) + \delta_{kt}$$

where Δ is the difference over time.

Taking the difference again with respect to the geometric average over the varieties eliminates the intercept, and yields:

$$\begin{aligned}
\Delta \ln \bar{s}_{kt} &= -(\sigma - 1) \Delta \ln(\bar{p}_{kt}) + \varepsilon_{kt} \\
\Delta \ln \bar{s}_{kt} &= \frac{\omega}{1 + \omega} \Delta \ln(\bar{p}_{kt}) + \delta_{kt}
\end{aligned}$$

with $\bar{s}_{kt} = \ln(\frac{s_{kt}}{\bar{s}_t})$. Exploiting the assumption made by Feenstra (1994) that the double differences demand and supply shocks are orthogonal and heteroskedastic Broda and Weinstein (2006) write the following equation:

$$\underbrace{(\Delta \ln(\bar{p}_{kt}))^2}_{Y_{kt}} = \underbrace{\frac{\omega}{(1+\omega)(\sigma-1)}}_{\Theta_1} \underbrace{(\Delta \ln(\bar{p}_{kt}))^2}_{X_{kt}^1} + \underbrace{\frac{1-\omega(\sigma-2)}{(1+\omega)(\sigma-1)}}_{\Theta_2} \underbrace{(\Delta \ln(\bar{p}_{kt}))^2}_{X_{kt}^2} + \underbrace{\varepsilon_{kt} * \delta_{kt}}_{u_{kt}}$$

Since u_{kt} is correlated with X_{kt}^1 and X_{kt}^2 , Feenstra (1994) proposes to take the average over time and estimate the equation by running weighted least squares. Once the thetas are known, one can solve for σ .²³ In a previous version of the paper, Redding and Weinstein (2019) used an estimator which combines the demand system with the expenditure function and then minimize the difference between the change in the cost of living using the preferences in t and $t-1$. With this method we find $\sigma = 2.8$ which is a bit lower, however does not change the findings from above.

D Optimal Rate of Inflation: Consumer Valuation Bias

To study the impact of consumer valuation bias on optimal inflation, we follow Schmitt-Grohé and Uribe (2011). In their paper, they consider the quality bias, which is very closely related to the consumer valuation bias.

- **Households:**

Households exhibit the following lifetime utility function

$$E_0 \sum_{t=0}^{\infty} \beta^t U(a_t, h_t)$$

where a_t is the quantity of the composite good a household wants to consume, and h_t is hours worked. The demand for good i can be derived from the cost-minimization

²³Note that Redding and Weinstein (2019) use a slightly different approach where they weight the data for each good by the number of raw buyers.

problem for the composite good and yields

$$q_{it} = \frac{a_t}{\varphi_{it}} (p_{it}/\varphi_{it})^{-\sigma} \left[\sum_{k \in \Omega_t} (p_{kt}/\varphi_{kt})^{1-\sigma} \right]^{\frac{\sigma}{1-\sigma}}$$

$$= \frac{a_t}{\varphi_{it}} \left[\frac{f_{it}}{F_t} \right]^{-\sigma}$$

with $f_{it} = p_{it}/\varphi_{it}$ and $F_t = \left[\sum_{k \in \Omega_t} (f_{kt})^{1-\sigma} \right]^{\frac{1}{1-\sigma}}$. φ_{it} are the preference shocks for good i .

- **Firms:**

Monopolistic competitive firms exhibit linear production functions with labor as the only input, i.e., $z_t h_{it}$, where z_t is an aggregate productivity shock. w_{t+j} is the nominal wage rate. Firms can change prices only with probability θ . Firms therefore maximize the expected present discounted value of profits

$$\max_{\tilde{p}_{it}} \sum_{j=0}^{\infty} E_t r_{t,t+j} \theta^j [\tilde{p}_{it} q_{it+j} - w_{t+j} h_{it+j}]$$

subject to

$$q_{it} = z_t h_{it}$$

where $r_{t,t+j}$ is a nominal stochastic discount factor, and \tilde{p}_{it} is the optimal price if a firm can change the price.

The FOC is:

$$\sum_{j=0}^{\infty} E_t r_{t,t+j} \theta^j \left[\frac{\partial q_{it+j}}{\partial \tilde{p}_{it}} \left(\tilde{p}_{it} - \frac{w_{t+j}}{z_{t+j}} \right) + q_{it+j} \right] = 0$$

Note that since preference shocks hit firms by surprise, $E_t(\varphi_{it+j}) = \bar{\varphi}$ for $j > 0$. Hence,

$$\begin{aligned} q_{it+j} &= \frac{a_{t+j}}{E_t(\varphi_{it+j})} \left[\frac{\frac{\tilde{p}_{it}}{E_t(\varphi_{it+j})}}{\left(\sum_{i \in \Omega_t} \left(\frac{\tilde{p}_{it}}{E_t(\varphi_{it+j})} \right)^{1-\sigma} \right)^{\frac{1}{1-\sigma}}} \right]^{-\sigma} \\ &= \frac{a_{t+j}}{\bar{\varphi}} \left[\frac{\tilde{p}_{it}}{\left(\sum_{i \in \Omega_t} (\tilde{p}_{it})^{1-\sigma} \right)^{\frac{1}{1-\sigma}}} \right]^{-\sigma} \end{aligned}$$

Thus,

$$\frac{\partial q_{it+j}}{\partial \tilde{p}_{it}} = -q_{it+j} \frac{\sigma}{\tilde{p}_{it}}$$

Plugging this back into the FOC gives

$$\begin{aligned} & \sum_{j=0}^{\infty} E_t r_{t,t+j} \theta^j \left[-q_{it+j} \frac{\sigma}{\tilde{p}_{it}} \left(\tilde{p}_{it} - \frac{w_{t+j}}{z_{t+j}} \right) + q_{it+j} \right] \\ &= \sum_{j=0}^{\infty} E_t r_{t,t+j} \theta^j \left[-q_{it+j} \sigma \left(1 - \frac{w_{t+j}}{\tilde{p}_{it} z_{t+j}} \right) + q_{it+j} \right] \\ &= \sum_{j=0}^{\infty} E_t r_{t,t+j} \theta^j \left[-q_{it+j} \left(\sigma - 1 - \frac{\sigma w_{t+j}}{\tilde{p}_{it} z_{t+j}} \right) \right] \\ &= \sum_{j=0}^{\infty} E_t r_{t,t+j} \theta^j \left[q_{it+j} \left(\frac{\sigma-1}{\sigma} \tilde{p}_{it} - \frac{w_{t+j}}{z_{t+j}} \right) \right] \\ &= \sum_{j=0}^{\infty} E_t r_{t,t+j} \theta^j \left[\frac{a_{t+j}}{\bar{\varphi}} \left[\frac{\tilde{p}_{it}}{\left(\sum_{i \in \Omega_t} (\tilde{p}_{it})^{1-\sigma} \right)^{\frac{1}{1-\sigma}}} \right]^{-\sigma} \left(\frac{\sigma-1}{\sigma} \tilde{p}_{it} - \frac{w_{t+j}}{z_{t+j}} \right) \right] = 0 \end{aligned}$$

This is the same expression as that in Schmitt-Grohé and Uribe (2011) p. 60; hence, we can follow their conclusions.