Universität Basel

Wirtschaftswissenschaftliche Fakultät

WWZ forum

January 2019

Sources of Bias in Inflation Rates and Implications for Inflation Dynamics

WWZ Working Paper 2019/02

Rahel Braun, Sarah M. Lein

A publication of the Center of Business and Economics (WWZ), University of Basel.

Universität Basel
Peter Merian-Weg 6
4052 Basel, Switzerland
wwz.unibas.ch

Corresponding Author:
Rahel Braun
Tel.: +41 61 207 07 83
Mail: rahel.braun@unibas.ch

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

Rahel Braun† and Sarah M. Lein‡

17 January, 2019

Abstract

Official statistics measuring the cost of living are known to suffer from several biases because they often do not accurately capture substitution patterns, product entry/exit, and preference shifts. In particular, the latter two biases have been shown to be large on average. This paper shows that the size of the biases can also vary with economic conditions. Using AC Nielsen homescan data for Switzerland around the 2015 appreciation of the CHF against the EUR, which resulted in a substantial shift in relative prices, it is first shown that official price indexes can be tracked using homescan data when applying the same methodology as that of the statistical office. Then, the often-acknowledged traditional substitution bias arising from lagged expenditure weights is shown to be relatively small. Based on an approach that allows for preference shocks and product entry/exit (unified price index), the two further biases, that is, the consumer valuation bias and the variety adjustment bias, are evaluated. Both are large on average and of similar size, together resulting in a 3.7 percentage point bias in the annual inflation rate. Furthermore, the bias is particularly large in the aftermath of the 2015 appreciation, increasing to 5.3 percentage points. In particular consumer valuation is shown to contribute significantly to the increase in the bias after the exchange-rate shock. The unified price index declines by 2.4 times more than the traditional Fisher index, suggesting that taking into account the time-variation in the bias is important.

JEL classification: E31, E4, E5, C3, C23
Keywords: Homescan data, relative price shifts, inflation measurement, bias in inflation indexes.

*We thank Ina Banik, Alfred Dijs, Petra Gerlach, Matthias Gubler, Christian Hepenstrick, Matthias Lutz, Stephen Redding, Barbara Rudolf, David Weinstein, participants of the SGVS annual meeting, the EEA-ESEM, the ESCoE Conference, the VfS annual meeting, the YSEM meeting, and seminar participants at the University of Basel and the Swiss National Bank for helpful comments and suggestions. We furthermore thank Maximilian Bastian and Lucienne Brunner for excellent research assistance. Funding for this project was generously provided by the WWZ Forum.

†Department of Business and Economics, University of Basel. Email: rahel.braun@unibas.ch

‡Department of Business and Economics, University of Basel, and KOF ETH Zurich. Email: sarah.lein@unibas.ch

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 new product 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 (2017) show that neglecting these consumer preference shifts can result in a potentially large bias, which they label the consumer valuation 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 how 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

[1]See also the Boskin Commission Report (Boskin et al., 1996).

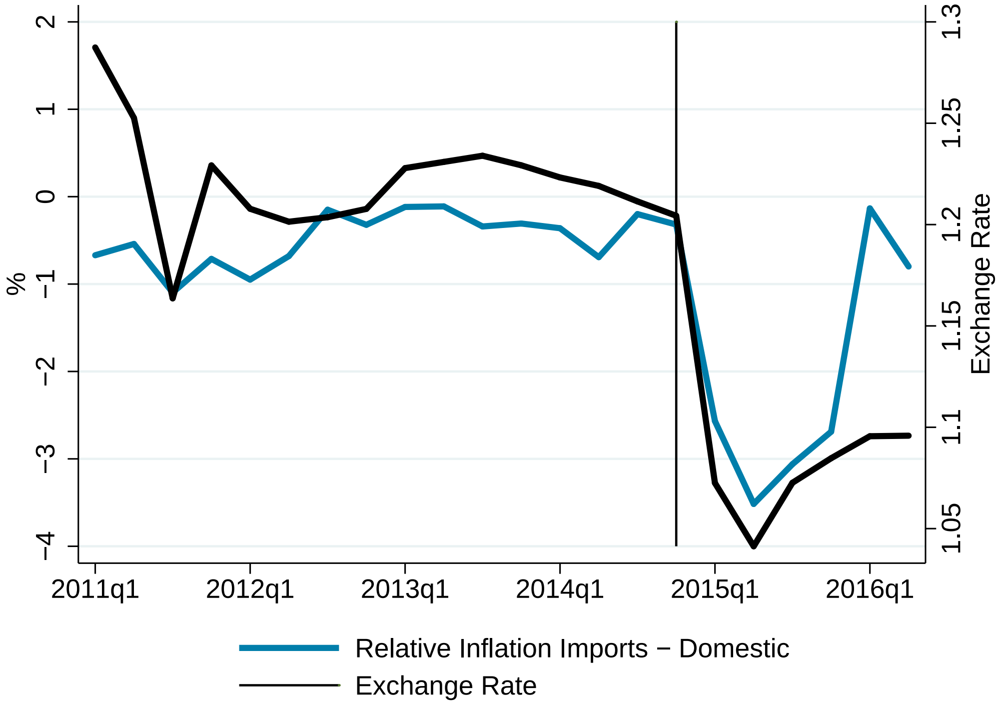
January 2015, as documented in Auer et al. (2017) and illustrated in Figure 1: consumer prices for imports declined by approximately 4 percent six months after the shock in response to the 14% appreciation of the CHF against the EUR, while domestically produced product's prices remained relatively stable.[2] As a result of the decline in relative prices for imports, consumers switch expenditures towards imported products (Auer et al., 2017). 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. The consumer valuation bias has a similar logic as that of the quality bias: prices should be adjusted for preferences, similarly to the commonly used adjustment for quality changes. If consumers' valuation for a product increases at an unchanged price, its price adjusted for preferences should decrease. Therefore, if relative demand shifts not only because prices change but also because consumers' preferences adjust, the consumer valuation bias may be larger in periods where relative prices change to a greater extent, depending on how much expenditures shift towards cheaper and more appealing products.

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 unified price index (UPI), which is not subject to traditional substitution bias, allows for product entry and exit, and captures consumer preference shifts. By observing the differences between the two indexes, we find similar average levels of bias compared to those of the US figures documented in Redding

[2]The 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 patters and relative prices. See Auer et al. (2017) for a detailed description of this unique episode.

[3]The 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.

and Weinstein (2017): neglecting the preference shifts results in a bias of 1.73 percentage points, and neglecting product entry/exit results in a bias of 2 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 net product entry bias, we therefore do not touch upon it in our analysis. The traditional substitution bias, which results from using lagged weights, is very small (<0.4 percentage points), also in line with previous literature for the US and other countries.[4]

The size of the bias increases by an additional 2.6 percentage points after the shock to relative prices. While the traditional substitution bias remains small in the period of the large shock, net product entry and preference shifts lead to a 2.4 times larger response of the UPI

[4]The 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 et al., 1996). For Switzerland, Diewert et al. (2009) estimate the annual substitution bias to be 0.13% on average between the years 1993 and 2002.

to the exchange-rate shock than what is suggested by the official price index. The largest contribution of this additional decline comes from preference shifts: even though consumers are faced with substantially low prices for imported products compared to domestically produced goods, expenditures for domestically produced goods remain high. If the relative prices of domestically produced goods increase but expenditures on these goods do not decline substantially, the UPI allocates these price and expenditure patterns to preference shifts. In other words, prices for domestically produced goods, adjusted for preferences, decline by more than prices weighted by expenditure shares in the official price indexes suggest.

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 (2017), 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 goods 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.

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 et al. (2017). 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

[5]Here we differ from Auer et al. (2017), 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.

[6]These 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 goods 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 goods 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.

Table 2: *Data overview*

	Transactions	Expenditure Share (%)	Products
Full Sample: Ω_t			
Total	13,856,430	100	116,799
Swiss	10,682,246	75.46	59,447
Imports	3,174,184	24.60	57,352
Imports EA	2,386,006	17.58	49,161
Imports ROW	788,178	7.04	8,191
Common Goods: $\Omega_{t-1,t}$			
Total	10,285,254	100	60,964
Swiss	7,945,897	75.86	35,674
Imports	2,339,357	24.14	25,290
Imports EA	1,744,744	17.10	21,289
Imports ROW	594,613	7.04	4,001
Super Common Goods: $\Omega_{\forall t}$			
Total	7,369,115	100	6,700
Swiss	5,782,661	76.75	4,689
Imports	1,586,454	23.25	2,011
Imports EA	1,166,400	16.33	1,509
Imports ROW	1,166,400	6.92	502

Notes: Transactions are the number of purchases observed, Expenditure Share is the share of total expenditures (in %), and Products shows the number of unique products in the respective sample. 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.

[7] The 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

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 *product* k means the unique product (EAN). For example, for the *productclass* rice, and the *products* 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 0.71), 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 is the expenditure for imports over total expenditure within a group.

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

	SFSO share %	*share* %	*import ratio* %
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 data set.

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 index. [8]

3 Price indexes

In this section, we first summarize the traditional Laspeyres price index and a slight modification, the Young index, which are widely used indexes in official CPI statistics. We include the Young index because it is the specific index used by the SFSO, to which we compare our data. Both indexes are based on lagged consumption expenditures, which are used to weight the current period prices. Since consumers tend to substitute more expensive for cheaper goods, these indexes overvalue 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 shifts in preferences and does not consider the utility gain from net product entry. The UPI developed in Redding and Weinstein (2017) overcomes these limitations: it is much more flexible and nests all the other indexes discussed above.

3.1 Laspeyres, Paasche, and Fisher Price Indexes

The formula for the Laspeyres (1871) index is

$$\begin{aligned} \Phi_t^{Las} &= \left[\sum_{j \in \Omega_{t,t-1}} S_{jt-1} \left(\frac{p_{jt}}{p_{jt-1}} \right) \right] \\ \text{with} \quad S_{jt-1} &= \frac{p_{jt-1} q_{jt-1}}{\sum_l p_{lt-1} q_{lt-1}}, \end{aligned}$$

where t-1 is any period in the past. j stands for product k or productclass n. p_{jt}/q_{jt} is price/quantity of good k purchased at time t. Because we first compare the information

[8] See Swiss Federal Statistical Office (2016).

included in the homescan data to the information in the CPI, we follow the exact method chosen by the SFSO. This method is similar to the Laspeyres index and is called the Young (1812) index. The formula is:

$$\Phi_t^{You} = \left[\sum_{j \in \Omega_{t,t-1}} S_{jb} \left(\frac{p_{jt}}{p_{jt-1}} \right) \right]$$
$$\text{with } ^9 \quad S_{jb} = \frac{p_{jb} q_{jb}}{\sum_l p_{lb} q_{lb}}. \tag{1}$$

The SFSO collects the weights in an annual expenditure survey,[10]. 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 $b = t - 8$ [11] in equation (1). 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 Paasche (1874) index uses weights from the current period t instead of a previous period t-1.

$$\begin{aligned} \Phi_t^{Paa} &= \left[\sum_{j \in \Omega_{t,t-1}} S_{jt} \left(\frac{p_{kt}}{p_{jt-1}} \right)^{-1} \right]^{-1} \\ \text{with} \quad S_{jt} &= \frac{p_{jt} q_{jt}}{\sum_l p_{lt} q_{lt}}. \end{aligned}$$

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,

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

[10] See Swiss Federal Statistical Office (2013)

[11] Since 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.

3.2 Unified Price Index

Recently, Redding and Weinstein (2017) 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 important roles and that including them lowers the cost of living substantially.

The UPI 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.[12] 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:

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

$$\text{with } \omega^*_{kt} \equiv \frac{\frac{S^*_{kt} - S^*_{kt-1}}{ln(S^*_{kt}) - ln(S^*_{kt-1})}}{\sum_{l \in \Omega_{t,t-1}} \frac{S^*_{lt} - S^*_{lt-1}}{ln(S^*_{lt}) - ln(S^*_{lt-1})}}$$

Redding and Weinstein (2017) expand this approach with two features. They also assume

[12] The inflation calculation with the economic approach and the resulting cost-of-living index go 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.

constant elasticity of substitution (CES) preferences [13]

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

however, they add preference (taste) parameters $\varphi_{kt} > 0$ and include new and exiting products (Ω_t) rather than only goods bought in both periods ($\Omega_{t,t-1}$).[14] Household optimization then yields the following unit ($U^0 = 1$) expenditure function [15]

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

and expenditure shares (applying Shepard's Lemma)

$$S_{kt} \equiv \frac{p_{kt} q_{kt}}{\sum\limits_{l \in \Omega_t} p_{lt} q_{lt}} = \frac{(p_{kt}/\varphi_{kt})^{1-\sigma}}{\sum\limits_{l \in \Omega_t} (p_{lt}/\varphi_{lt})^{1-\sigma}}. \tag{2}$$

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

$$\Phi^{CES}_{t-1,t} = \frac{\mathbb{P}_t}{\mathbb{P}_{t-1}} = \left[\frac{\sum\limits_{k \in \Omega_t} (p_{kt}/\varphi_{kt})^{1-\sigma}}{\sum\limits_{k \in \Omega_{t-1}} (p_{kt-1}/\varphi_{kt-1})^{1-\sigma}} \right]^{\frac{1}{1-\sigma}} \tag{3}$$

To account for the role of product entry and exit, Redding and Weinstein (2017) decompose equation (3) into a unit expenditure function on common goods 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 goods) as a share of total expenditures in t

[13]The authors show that the main results also hold for other preference functions.

[14]The authors base these results on Feenstra (1994), who originally demonstrated how to incorporate new and exiting goods.

[15]see Appendix B for the derivation

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

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

Similarly,

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

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

Hence, equation (3) can be written as

$$\begin{aligned} \Phi^{CES}_{t-1,t} &= \frac{\mathbb{P}_t}{\mathbb{P}_{t-1}} \\ &= \left[\frac{\lambda_{t-1}}{\lambda_t} \frac{\sum\limits_{k \in \Omega_{t-1,t}} \left(\frac{p_{kt}}{\varphi_{kt}}\right)^{1-\sigma}}{\sum\limits_{k \in \Omega_{t-1,t}} \left(\frac{P_{kt-1}}{\varphi_{kt-1}}\right)^{1-\sigma}} \right]^{\frac{1}{1-\sigma}} \\ &= \left(\frac{\lambda_{t-1}}{\lambda_t}\right)^{\frac{1}{1-\sigma}} \frac{\mathbb{P}^*_t}{\mathbb{P}^*_{t-1}} \end{aligned} \tag{4}$$

where $\mathbb{P}^*_t \equiv \left[\sum\limits_{k \in \Omega_{t-1,t}} \left(\frac{P_{kt}}{\varphi_{kt}}\right)^{1-\sigma} \right]^{\frac{1}{1-\sigma}}$ is the unit expenditure function defined over the set of common goods $\Omega_{t-1,t}$.

To make the price index money-metric, Redding and Weinstein (2017) assume a constant geometric mean of all demand parameters $\bar{\varphi} \Rightarrow \frac{1}{N_{t,t-1}} \sum\limits_{k \in \Omega_{t,t-1}} (ln(\varphi_{kt}) - ln(\varphi_{kt-1})) = 0.$

Rewriting and combining equation (4) yields the UPI [16],[17]:

$$\Phi_{t-1,t}^{UPI} = \underbrace{\left(\frac{\lambda_t}{\lambda_{t-1}}\right)^{\frac{1}{\sigma-1}}}_{\text{Variety Adjustment}} \left[\underbrace{\frac{\tilde{P}_t^*}{\tilde{P}_{t-1}^*}}_{\text{Jevons}} \underbrace{\left(\frac{\tilde{S}_t^*}{\tilde{S}_{t-1}^*}\right)^{\frac{1}{\sigma-1}}}_{\text{Expenditure Shifts}}\right] \tag{5}$$

where a tilde represents the geometric average and an asterisk denotes the set of common goods 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 goods available in t and $t-1$.

Equation (5) shows that the UPI is the product of three terms: the variety adjustment term, the ratio of two Jevons price indexes for common goods at time t and $t-1$, and the expenditure shifts term for common goods. 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 goods, $\tilde{S}_t^*$. If this term decreases, that is $\tilde{S}_t^* < \tilde{S}_{t-1}^*$, expenditure shares become more uneven across products. 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.[18]

The estimation of the elasticity of substitution and preference parameters is based on the double-reverse weighting estimation procedures used by Redding and Weinstein (2017), which

[16]For the detailed derivation see (Redding and Weinstein, 2017) and their web appendix.

[17]It is called the *unified* price index because it unifies the observed shifts in expenditure shares from micro data with macro money-metric utility functions, which makes it possible to compare welfare over time and nests many commonly used price indexes.

[18]This expenditure shifts term is only negligible if either the elasticity of substitution is infinity, that is, goods are prefect substitutes, or if all preference-adjusted prices are identical and therefore all expenditure shares are equal across products. When preference-adjusted prices become more dispersed, so do expenditure shares across products. 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).

are outlined in Appendix C. Our estimates of the elasticity of substitution yields a $\sigma = 2.8$ on average, close to estimates of demand elasticity reported in the literature using similar micro data.[19]

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 price index applying the Fisher CPI-method deviates from that of the UPI, where the largest contribution to these differences come from, and whether the contributions change after the price shock.

Figure 2

(a) Homescan and CPI (b) Homescan indexes

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.

4.1 Comparison with the official index

Panel (a) of Figure 2 shows the CPI calculated with the *super common goods* homescan dataset ($\Omega_{\forall t}$) and the official CPI for food, beverages and tobacco. The inflation rate is then calculated as the yearly log difference. The official CPI figures (we always compare with the index for the same product categories not the overall index including services, etc.) are very

[19]As we estimate one elasticity of substitution for each time period, we also considered using only the elasticity estimates from the year 2015, where we have more exogenous variation in prices due to the nominal cost shock, but the estimates from 2015 Q1 to Q4 were very close to this average (Q1: 2.3, Q2: 2.3, Q3: 2.5, Q4: 2.8), such that the overall results are not affected by using the 2015-estimates for the elasticity of substitution.

close to those of 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.

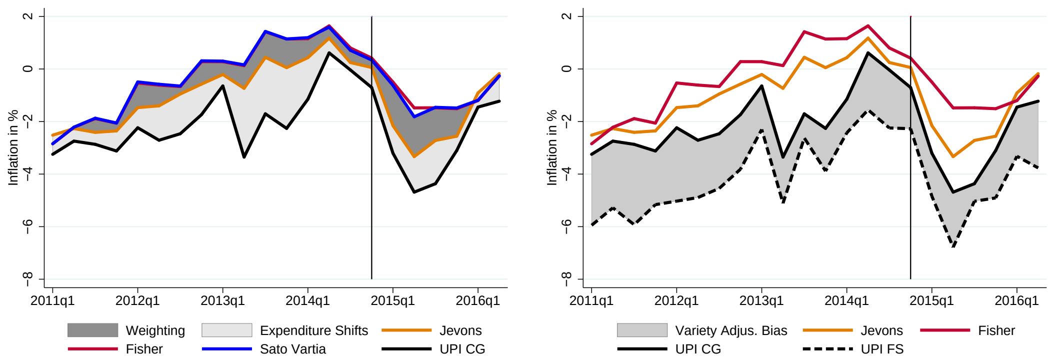
4.2 Bias decomposition

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 size of the traditional substitution bias. To calculate the difference, we focus on the common goods 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 0.45 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 et al. (1996), Kaplan and Schulhofer-Wohl (2017)) and slightly higher than the one for Switzerland (Diewert et al., 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.

Consumer valuation bias. Having shown that the traditional substitution bias is small, we examine the impact of omitting the preference parameters. Panel (a) of Figure 3 plots the Fisher index in red, the Jevons index in orange, and the UPI-CG (unified price index based on common-goods sample) in black.[20] The sum of the two gray areas is the total bias arising from neglecting the preference parameters and amounts to 1.73 percentage points. The bias can be decomposed into two parts. The first 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 products. This difference is 0.64 percentage points on average. The second part is the shift between the Jevons index and the UPI-CG, which amounts to 1.09 percentage points.

Figure 3: *The role of consumer valuation and net product entry*



(a) Consumer valuation (b) Net product entry

This figure shows the Weightung and Expenditure Shifts Bias (Panel (a)) as well as the Variety Adjustment Bias (Panel (b)) together with Jevons, Fisher, Sato-Vartia and Unified Price Index. CG stands for the common goods sample while FS stands for the full sample as defined in Table (2).

Net product entry bias. Finally, we quantify the role of net product entry by including the variety adjustment term from Formula (5). The bias arising from neglecting this term is substantial. Panel (b) of Figure 3 shows that controlling for net product entry shifts inflation down by 1.97 percentage points, which is a large and persistent bias.

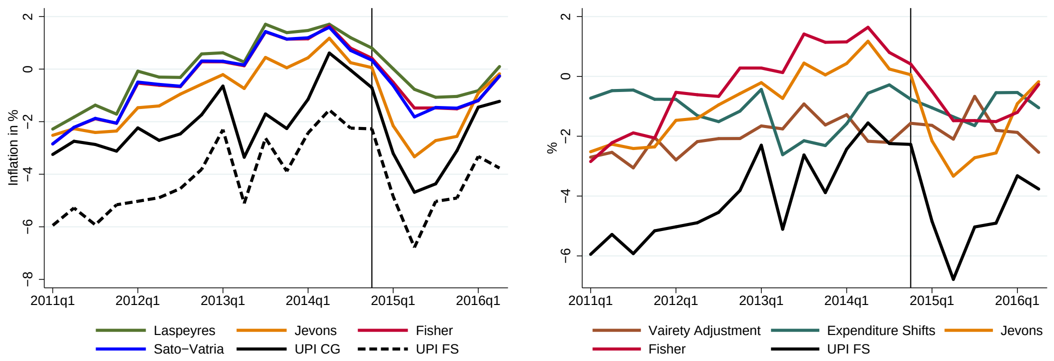
4.3 The role of biases in inflation dynamics

We have shown that, on average, neglecting preference shifts results in a bias of 1.73 percentage points and neglecting product entry/exit results in a bias of 1.97 percentage points. In this section, study whether the biases differ after relative price changes. Therefore, in Panel

[20]Note that the Sato-Vartia index, which is calculated with the economic approach, and the Fisher index are almost the same.

(a) of Figure 4, we plot all the previous indexes again and focus on the time after the exchange rate shock in January 2015.

Figure 4



(a) Dynamics (b) Decomposition

This figure shows the Laspeyres, Jevons, Fisher, Sato-Vartia and Unified Price Index, as well as the Variety Adjustment and Expenditure Shifts Bias defined below in equation (7). CG stands for common goods sample while FS stands for full sample as defined in Table (2).

The indexes including the preference shocks (UPI) clearly decline much more than the indexes not including them (Fisher, Sato-Vartia). The UPI for the full sample declines by 4.51 percentage points between 2014Q4 and 2015Q2, while the Sato-Vartia index declines by only 1.89 percentage points. In the following, we study why the UPI reacts more to the shock. To do so, we decompose the UPI into its three terms and plot them separately (Panel (b) of Figure 4).

$$ln(\Phi^{UPI}_{t-1,t}) = \underbrace{\frac{1}{\sigma-1} * ln\left(\frac{\lambda_t}{\lambda_{t-1}}\right)}_{\text{Variety Adjustment}} + \underbrace{ln\left(\frac{\tilde{P}^*_t}{\tilde{P}^*_{t-1}}\right)}_{\text{Jevons}} + \underbrace{\frac{1}{\sigma-1} * ln\left(\frac{\tilde{S}^*_t}{\tilde{S}^*_{t-1}}\right)}_{\text{Expenditure Shifts}} \quad (6)$$

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

$$
\begin{aligned}
ln(\Phi_{t-1,t}^{UPI}) - ln(\Phi_{t-1,t}^{SV}) \quad = \quad & \underbrace{\overbrace{\frac{1}{\sigma-1} * ln\left(\frac{\tilde{S}_t^*}{\tilde{S}_{t-1}^*}\right)}^{\text{Expenditure Shifts}} + \overbrace{\sum_{k\in\Omega_{t,t-1}}\left(\frac{1}{N_{t,t-1}} - \omega_{kt}^*\right) ln\left(\frac{p_{kt}}{p_{kt-1}}\right)}^{Weightung}}_{ConsumerValuation} \\
& + \underbrace{\frac{1}{\sigma-1} * ln\left(\frac{\lambda_t}{\lambda_{t-1}}\right)}_{\text{Variety Adjustment}},
\end{aligned}
$$

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 expenditure shifts (explained above) 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).

The variety adjustment term and the expenditure shifts term do not react to the shock. However, the Jevons index (and hence the weighting term) reacts strongly. As the Jevons index uses uniform weights, this suggests that products with larger price declines tend to be those with relatively smaller expenditure weights in the Sato-Varita index. In this 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 goods sample, 40% of the products are imports, but they account for only approximately 23% of expenditures (Table 2). This explains why the Sato-Varita (or the Fisher index) decline by less: import prices decline by more but account for only 23%, compared to 40% for the Jevons index.

In the UPI, products are not weighted by expenditure shares, as in the Fisher 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 75% in our data and the official CPI), whose prices hardly change in response to the shock, mutes the response of the official Sato-Varita/Fisher-type price index to the exchange-rate appreciation. In the UPI, 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). Here, we compare to the Fisher index, because the Fisher index as been used to quantify the traditional substitution bias above. As shown in 4, so the same conclusions apply if we defined the bias as the difference between UPI and Sato-Vartia. Before the shock, the Fisher and the UPI differ by 3.72 percentage points on average (the standard deviation over time is 0.80). Almost half of the difference results from the neglected preference shocks (consumer valuation), and the other half (55%) results from variety adjustment. However, this changes after the shock. The difference between the Fisher and the UPI increases significantly, by 2.62 percentage points, and 80% of this difference comes from consumer valuation, while only 20% comes from variety adjustment.

Table 4: *Bias decomposition*

		2011-2014	Δ2014Q4-2015Q2
Fisher-UPI		3.72 (0.80)	2.62
shares			
Consumer Valuation		0.45	0.80
	Weighting	0.15	0.57
	Expenditure Shifts	0.30	0.23
Variety Adjustment		0.55	0.20

Notes: The first line shows the percentage point difference between the Fisher index and the UPI, on average, for 2011-2014 (first column) and between 2014Q4 and 2015Q2 (second column). The shares are calculated as $1 = \frac{Weighting}{Fisher-UPI} - \frac{ExpenditureShifts}{Fisher-UPI} - \frac{VarietyAdjustment}{Fisher-UPI}$. The number in parentheses in the line Fisher-UPI 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 (2017), neglecting preference shifts and the entry/exit of goods results in a 3.74 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 (2010) convincingly show that the optimal inflation rate might still be zero, even in the presence of unmeasured quality improvements. In Appendix D, we show that the preference parameters in the derivation of the UPI can also be interpreted as the quality adjustment term in Schmitt-Grohé and Uribe (2010). 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 (2010) 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 UPI, 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 et al. (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.[21] If statistical offices do not 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 shifts, 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 0.38 percentage points per year, whereas the consumer valuation bias is approximately 1.73 percentage points per year and the product entry bias is approximately 1.97 percentage points per year. These biases are substantial and in line with findings reported in Redding and Weinstein (2017) 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,

[21] In 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 et al., 2014).

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 UPI-inflation rate, which takes into account all biases, declines by 4.51 percentage points six months after the exchange rate shock, while the traditional Fisher inflation rate declines by only 1.89 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 UPI declines 2.4 times more than does the traditional official price 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 Fisher index and the UPI using a simple back-of-the-envelope calculation would result in a much larger increase in real private consumption after the appreciation (5.9% in 2015 and 4.4% in 2014).[22]

[22]Because the consumption deflator includes more product categories than those in our Fisher and UPI indexes, we calculate the growth rate of the consumption deflator and subtract the differences between the Fisher index and UPI 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.

References

Auer, R., A. Burstein, and S. M. Lein (2017): "Prices and Exchange Rates: Evidence from the 2015 Swiss Franc Appreciation," mimeo, UCLA.

Bilbiie, F. O., I. Fujiwara, and F. Ghironi (2014): "Optimal monetary policy with endogenous entry and product variety," *Journal of Monetary Economics*, 64, 1 – 20.

Boskin, M., D. Jorgenson, E. R. Dulberger, R. J. Gordon, and Z. Griliches (1996): "Final Report of the Advisory Commission to Study the Consumer Price Index," .

Cavallo, A. (2013): "Online and Official Price Indexes: Measuring Argentina's Inflation," *Journal of Monetary Economics*, 60, 152–65.

Diewert, W. (1976): "Exact and superlative index numbers," *Journal of Econometrics*, 4, 115–145.

Diewert, W. E., M. Huwiler, and U. Kohli (2009): "Retrospective Price Indices and Substitution Bias," *Swiss Journal of Economics and Statistics (SJES)*, 145, 127–135.

Feenstra, R. C. (1994): "New Product Varieties and the Measurement of International Prices," *The American Economic Review*, 84, 157–177.

Hausman, J. (2003): "Source of Bias and Solutions to Bias in the Consumer Price Index," 17, 23–44.

Kaplan, G. and S. Schulhofer-Wohl (2017): "Inflation at the household level," *Journal of Monetary Economics*, 91, 19 – 38, the Swiss National Bank/Study Center Gerzensee special issue title is "A Conference in Honor of Robert King".

Konüs, A. A. (1924): "The Problem of the True Index of the Cost of Living," *The Economic Bulletin of the Institute of Economic Conjuncture*, translated in English in 1939 in Econometrica,.

Laspeyres, E. (1871): "Die Berechnung einer mittleren Waarenpreissteigerung," *Jahrbücher für Nationalökonomie und Statistik*, 16, 296–314.

LLOYD, P. J. (1975): "Substitution Effects and Biases in Nontrue Price Indices," *American Economic Review*, 65, 301–13.

MOULTON, B. R. (1975): "Constant Elasticity Cost-of-Living Index in Share Relative Form," Washington, D.C.: Bureau of Labor Statistics).

NEARY, J. P. (2004): "Rationalizing the Penn World Table: True Multilateral Indices for International Comparisons of Real Income," *American Economic Review*, 94, 1411–1428.

PAASCHE, H. (1874): "Über die Preisentwicklung der letzten Jahre nach den Hamburger Börsennotirungen," *Jahrbücher für Nationalökonomie und Statistik*, 12, 168–178.

REDDING, S. J. AND D. E. WEINSTEIN (2017): "A Unified Approach to Estimating Demand and Welfare," Working Paper 22479, National Bureau of Economic Research.

——— (2018): "Accounting for Trade Patterns," NBER Working Papers 24051, National Bureau of Economic Research, Inc.

SATO, K. (1976): "The Ideal Log-Change Index Number," *The Review of Economics and Statistics*, 58, 223–28.

SCHMITT-GROHÉ, S. AND M. URIBE (2010): "The Optimal Rate of Inflation," NBER Working Papers 16054, National Bureau of Economic Research, Inc.

SWISS FEDERAL STATISTICAL OFFICE (2013): *Haushaltsbudgeterhebung 2011-Kommentierte Ergebnisse und Tabellen*, Swiss Federal Statistical Office, Neuchâtel.

——— (2016): *Landesindex der Konsumentenpreise: Dezember 2015 = 100 Methodenübersicht und Gewichtung 2016*, Swiss Federal Statistical Office, Neuchâtel.

VARTIA, Y. O. (1976): "Ideal Log-Change Index Numbers," *Scandinavian Journal of Statistics*, 3, 121–126.

YOUNG, A. (1812): "An Inquiry into the Progressive Value of Money in England as Marked by the Price of Agricultural Products," *Piccadilly: Hatchard*.

Appendix

A Inflation Rates with offical Prices from the Swiss Federal Statistical Office

Instead of taking the prices from the scanner data set we can also compute inflation rates with the prices collected by the Swiss Federal Statistical Office. Since data is not available on a barcode level but only for productclasses (e.g. rice) inflation rates are

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

$$\text{with } ^{23} \quad S_{jb} = \frac{p_{jb} q_{jb}}{\sum_l p_{lb} q_{lb}}.$$

B Derivation of the Unified Price Index

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

Budget constraint

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

The cost minimizing problem is:

$$\min_{q_{kt}} \sum_{k \in \Omega_t} p_{kt} q_{kt}$$

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

First order conditions:

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

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

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

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 and preference parameters

To calculate the UPI, it is necessary to first estimate the elasticity of substitution σ and the preference parameters φ.

To do so, Redding and Weinstein (2017) decompose equation (3) in two ways

1.

$$\frac{\mathbb{P}_t}{\mathbb{P}_{t-1}} = \left(\frac{\lambda_{t-1}}{\lambda_t}\right)^{\frac{1}{\sigma-1}} \Theta^F_{t-1,t} \left[\sum_{k\in\Omega_{t,t-1}} S^*_{kt-1}\left(\frac{P_{kt}}{P_{kt-1}}\right)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}$$

with

$$\Theta^F_{t,t-1} = \left[\frac{\sum_{k\in\Omega t,t-1} S_{kt-1}\left(\frac{P_{kt}}{P_{kt-1}}\right)^{1-\sigma}\left(\frac{\varphi_{kt}}{\varphi_{kt-1}}\right)^{1-\sigma}}{\sum_{k\in\Omega t,t-1} S_{kt-1}\left(\frac{P_{kt}}{P_{kt-1}}\right)^{1-\sigma}}\right]^{(\frac{1}{\sigma-1})}$$

2.

$$\frac{\mathbb{P}_t}{\mathbb{P}_{t-1}} = \left(\frac{\lambda_{t-1}}{\lambda_t}\right)^{\frac{1}{\sigma-1}} (\Theta^B_{t-1,t})^{-1} \left[\sum_{k\in\Omega_{t,t-1}} S^*_{kt}\left(\frac{P_{kt}}{P_{kt-1}}\right)^{-(1-\sigma)}\right]^{-\frac{1}{1-\sigma}}$$

with

$$\Theta^B_{t,t-1} = \left[\frac{\sum_{k\in\Omega t,t-1} S^*_{kt}\left(\frac{P_{kt-1}}{P_{kt}}\right)^{1-\sigma}\left(\frac{\varphi_{kt-1}}{\varphi_{kt}}\right)^{(\sigma-1)}}{\sum_{k\in\Omega t,t-1} S^*_{kt}\left(\frac{P_{kt-1}}{P_{kt}}\right)^{1-\sigma}}\right]^{\frac{1}{1-\sigma}}$$

They also make the identifying assumptions that

$$\Theta^F_{t,t-1} = (\Theta^B_{t,t-1})^{-1} = 1 \tag{7}$$

The preference parameters φ are replaced by the shares from equation (2)

$$\frac{\varphi_{kt}}{\bar{\varphi}} = \frac{P_{kt}}{\tilde{P}_t}\left(\frac{S_{kt}}{\tilde{S}_t}\right)^{\frac{1}{\sigma-1}} \tag{8}$$

which, together with assumption (7), yields the following two moment conditions.

$$\Theta^F_{t,t-1} - 1 = \left[\frac{\sum_{k\in\Omega t,t-1} S^*_{kt-1} \left(\frac{P_{kt}}{P_{kt-1}}\right)^{1-\sigma} \left(\frac{\frac{P_{kt}}{\tilde{P}_t}\left[\frac{S_{kt}}{\tilde{S}_t}\right]^{\frac{1}{\sigma-1}}}{\frac{P_{kt-1}}{\tilde{P}_{t-1}}\left[\frac{S_{kt-1}}{\tilde{S}_{t-1}}\right]^{\frac{1}{\sigma-1}}} \right)^{(\sigma-1)}}{\sum_{k\in\Omega t,t-1} S^*_{kt-1} \left(\frac{P_{kt}}{P_{kt-1}}\right)^{1-\sigma}} \right]^{(\frac{1}{1-\sigma})} - 1 = 0 \quad (9)$$

$$(\Theta^B_{t,t-1})^{-1} - 1 = \left[\frac{\sum_{k\in\Omega t,t-1} S^*_{kt} \left(\frac{P_{kt-1}}{P_{kt}}\right)^{1-\sigma} \left(\frac{\frac{P_{kt-1}}{\tilde{P}_{t-1}}\left[\frac{S_{kt-1}}{\tilde{S}_{t-1}}\right]^{\frac{1}{\sigma-1}}}{\frac{P_{kt}}{\tilde{P}_t}\left[\frac{S_{kt}}{\tilde{S}_t}\right]^{\frac{1}{\sigma-1}}} \right)^{(\sigma-1)}}{\sum_{k\in\Omega t,t-1} S^*_{kt} \left(\frac{P_{kt-1}}{P_{kt}}\right)^{1-\sigma}} \right]^{-(\frac{1}{1-\sigma})} - 1 = 0 \quad (10)$$

Since there is only one unknown σ_{RW} [24], it can be estimated with a general method of moments estimator (GMM).

Redding and Weinstein (2017) additionally prove several results.

First, σ_{RW} consistently estimates the true σ if

- the demand shocks are small $\left(\frac{\varphi_{kt}}{\varphi_{kt-1}} \to 1\right)$
- the number of common goods becomes large $\left(N_{t,t-1} \to \infty, \frac{\varphi_{kt}}{\varphi_{kt-1}} \text{ are i.i.d. }\right)$

Second, σ_{RW} depends on the correlation between price and demand shocks. If

- $cov\left(\left(\frac{\varphi_{kt}}{\varphi_{kt-1}}\right), \left(\frac{P_{kt}}{P_{kt-1}}\right)\right) > 0$

  $\Theta^F_{t-1,t} > 1 > \Theta^B_{t,t-1}$

  $\sigma > \sigma_{RW}$
- $cov\left(\left(\frac{\varphi_{kt}}{\varphi_{kt-1}}\right), \left(\frac{P_{kt}}{P_{kt-1}}\right)\right) < 0$

  $\Theta^B_{t,t-1} > 1 > \Theta^F_{t-1,t}$

  $\sigma < \sigma_{RW}$

The true σ is not identified because, depending on the correlation between price and demand

[24] RW stands for reverse weighting because "it involves equating expressions for the change in the cost of living using both initial-period and final-period expenditure share weights" (Redding and Weinstein, 2017, p.26)

shocks, σ can be below or above σ_{RW}. Therefore, the authors construct an estimator with the opposite interpretation; that is, they use the inverse demand shocks in equation (9) and (10) and estimate σ_{DRW} [25] with the same procedure as above. σ_{DRW} now has the opposite interpretation, such that

- if $cov\left(\left(\frac{\varphi_{kt}}{\varphi_{kt-1}}\right), \left(\frac{P_{kt}}{P_{kt-1}}\right)\right) > 0$ then $\sigma_{DRW} > \sigma > \sigma_{RW}$
- if $cov\left(\left(\frac{\varphi_{kt}}{\varphi_{kt-1}}\right), \left(\frac{P_{kt}}{P_{kt-1}}\right)\right) < 0$ then $\sigma_{DRW} < \sigma < \sigma_{RW}$.

In other words, it is possible to construct an upper and lower bound for the true elasticity of substitution (given that preferences are homothetic and the geometric average of the preference is constant across time).

Once σ is known, one can easily calculate the preference parameters φ_{kt} with equation (8).

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 (2010). 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 problem for the composite good and yields

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

[25] DWR stands for double reverse weighting. This approach comes from the literature of reverse regression.

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

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

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

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

This is the same expression as that in Schmitt-Grohé and Uribe (2010) p. 60; hence, we can follow their conclusions.