Estimation of the Coffee Price Index Using Scanner Data: the Choice of the Micro Index

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Abstract: With the use of Nielsen scanner data on coffee sales Statistics Netherlands is undertaking empirical research into the effects of alternative micro indexes and different commodity sampling designs on the (Laspeyres-type) consumer price subindex of coffee. In this paper we are concerned with the choice of the micro index formula. We also analyse the effect of the frequency of price observation.

1. Introduction

Statistics Netherlands is undertaking empirical research using scanner data from A.C. Nielsen (Nederland) B.V. to gain insight into the effects of alternative micro index formulas and sampling designs on the Laspeyres-type consumer price subindex of coffee. It is hoped that the findings may, to some extent, be generalised to other commodity groups. De Haan and Opperdoes (1997) demonstrate how current CPI practices can be approximated using the Nielsen data. In this paper we study the effect of the choice of the micro index, that is the price index formula at the lowest aggregation level. In addition to formulas that use price data only (such as the ratio of unweighted average prices which is used for the official coffee price index and thus already calculated in the simulation study just mentioned), the Laspeyres, Paasche and Fisher price indexes are calculated as well as the unit value index. We will also be looking at the effect of sampling in time. The usual single price quotation is approximated by the unit value over the observation week, while the unit value taken over the whole month reflects the average transaction price.

Section 2 of this paper gives an overview of the various micro price indexes that we will calculate. Section 3 describes the unit value index. Section 4 goes into sampling in time. Section 5 presents empirical results, and section 6 concludes.

2. The price indexes at the elementary aggregation level

2a. Estimating a fixed weight commodity group price index

Let commodity group *A* consist of a finite number of commodities (items); $g \in A$ means that *g* belongs to *A*. We assume that *A* is fixed during time. The Laspeyres-type (fixed weight) price index of commodity group *A* in period *t* is

(1) $$P^t = \sum_{g \in A} w_g^0 P_g^t ,$$

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where P_g^t is the price index of item *g* and w_g^0 is the base period expenditure share of *g* within *A*. Let $\hat{A}$ denote a sample of items taken from *A*. The item group price index is estimated by

(2) $$\hat{P}^t = \sum_{g \in \hat{A}} \hat{w}_g^0 P_g^t ,$$

where $\hat{w}_g^0$ is the expenditure share of the sampled item *g* with respect to the entire sample. Let B_g denote the set of outlets in which commodity *g* can be bought; $b \in B_g$ means that *g* can be bought in outlet *b*. Usually B_g is subdivided into strata B_{gi} (numbered *i= 1,...,I*) with size N_{gi}. We assume that the strata are fixed during time. The Laspeyres-type price index of item *g* becomes

(3) $$P_g^t = \sum_{i=1}^{I} w_{gi}^0 P_{gi}^t$$

where P_{gi}^t is the price index of stratum *i*, also called the micro price index (and sometimes referred to as the elementary aggregate price index). The stratum weights reflect relative base period expenditures. All $b \in B_{gi}$ are supposed to be observed, so that we abstain from the sampling of outlets (sampling in space).

2b. Micro indexes based on price data only

Outlet-specific quantity and/or expenditure data are hard to come by, at least until Nielsen data became available. In practice stratum price indexes are still calculated with the use of price data only. Statistics Netherlands currently uses the *ratio of (unweighted) average prices*

(4) $$P_{gi}^t(R) = \frac{\sum_{b \in B_{gi}} p_{gb}^t \big/ N_{gi}}{\sum_{b \in B_{gi}} p_{gb}^0 \big/ N_{gi}} ,$$

where p_{gb}^s is the price of item *g* at outlet *b* in period *s* (s=0, *t*). An alternative would be the (*unweighted) arithmetic mean of price relatives*

(5) $$P_{gi}^t(A) = \sum_{b \in B_{gi}} (p_{gb}^t \big/ p_{gb}^0) \big/ N_{gi} .$$

Instead of the arithmetic mean one could also choose the *geometric mean of price relatives*

(6) $$P_{gi}^t(G) = \prod_{b \in B_{gi}} (p_{gb}^t \big/ p_{gb}^0)^{1/N_{gi}} .$$

Notice that this is identical to the ratio of geometric averages of prices. From the well-known relation between the arithmetic and the geometric mean of positive real numbers it follows that

$P_{gi}^{t}(G) \leq P_{gi}^{t}(A)$. The Fourteenth International Conference of Labour Statisticians (see Turvey, 1989) recommended that "In the calculation of elementary aggregate indices, consideration should be given to the possible use of geometric means". Some European countries do actually use them for their CPIs. In the USA the Advisory Commission to Study the Consumer Price Index recommended that the Bureau of Labor Statistics "should change its procedure for combining price quotations by moving to geometric means at the elementary aggregation level" (Boskin et al., 1996).

2c. Micro indexes based on price and quantity data

With the availability of quantity and/or expenditure data for individual outlets we are able to calculate the fixed weight or *Laspeyres price index*

$$(7) \qquad P_{gi}^{t}(L) = \frac{\sum_{b \in B_{gi}} p_{gb}^{t} x_{gb}^{0}}{\sum_{b \in B_{gi}} p_{gb}^{0} x_{gb}^{0}} = \sum_{b \in B_{gi}} w_{gb}^{0} (p_{gb}^{t} / p_{gb}^{0}),$$

where x_{gb}^{0} denotes the quantity of *g* bought at outlet *b* during the base period and w_{gb}^{0} the corresponding turnover share. Balk (1994) points out that under specific assumptions the Laspeyres stratum index reduces to expressions (4) or (5). See also the Appendix.

If we abandon the fixed weight concept within strata, it may be worthwile to look at the *Paasche price index*

$$(8) \qquad P_{gi}^{t}(P) = \frac{\sum_{b \in B_{gi}} p_{gb}^{t} x_{gb}^{t}}{\sum_{b \in B_{gi}} p_{gb}^{0} x_{gb}^{t}},$$

where x_{gb}^{t} is the quantity of *g* bought at *b* in period *t*. However, when allowing substitution of item sales between outlets within strata a better solution would probably be to construct the *Fisher price index*

$$(9) \qquad P_{gi}^{t}(F) = \sqrt{P_{gi}^{t}(L) P_{gi}^{t}(P)}.$$

2d. Chained indexes

In the foregoing, the prices in period t were compared directly to the base period prices. In addition to the above mentioned direct price indexes, we will compute the monthly chained equivalents

$$(10) \qquad P_{gi}^{t}(.,c) = \prod_{\tau=1}^{t} P_{gi}^{\tau/\tau-1}(.),$$

where $P_{gi}^{\tau/\tau-1}(.)$ denotes the (direct) price index of period τ relative to period τ -1. It is easy to verify that the direct and chained indexes coincide in case of the ratio of average prices and the geometric mean of price relatives.

3. The unit value index

There is another class of indexes where substitution between outlets comes into play. To start with, rewrite expression (7) as

$$(7') \qquad P_{gi}^{t}(L) = \frac{\sum_{b \in B_{gi}} z_{gb}^{0} p_{gb}^{t}}{\sum_{b \in B_{gi}} z_{gb}^{0} p_{gb}^{0}},$$

where $z_{gb}^{0} = x_{gb}^{0} \Big/ \sum_{b' \in B_{gi}} x_{gb'}^{0}$.

In words: the Laspeyres micro index is a ratio of weighted average prices with weights reflecting relative base year quantities. But why should we weight prices in period *t* by base period quantities to obtain the average price? Why not weight them by quantities in period *t*? By doing so, we are constructing the *unit value index*

$$(11) \qquad P_{gi}^{t}(U) = \frac{\sum_{b \in B_{gi}} z_{gb}^{t} p_{gb}^{t}}{\sum_{b \in B_{gi}} z_{gb}^{0} p_{gb}^{0}} = \frac{\sum_{b \in B_{gi}} v_{gb}^{t} \Big/ \sum_{b \in B_{gi}} x_{gb}^{t}}{\sum_{b \in B_{gi}} v_{gb}^{0} \Big/ \sum_{b \in B_{gi}} x_{gb}^{0}},$$

where $z_{gb}^{s} = x_{gb}^{s} \Big/ \sum_{b' \in B_{gi}} x_{gb'}^{s}$ and $v_{gb}^{s} = p_{gb}^{s} x_{gb}^{s}$ (s=0,t).

According to the axiomatic theory of index numbers the unit value index cannot be called a price index, since it satisfies neither the dimensionality axiom nor the proportionality axiom (Balk, 1995). The first axiom states that a price index should be insensitive to the units of measurement. The second axiom says that if all prices change by the same factor then the price index should be equal to this factor. That the unit value index does not pass the proportionality axiom also means that if the prices in period *t* are equal to the base period prices, the unit value index may show up with a value different from one (it fails the identity test), something that price statisticians probably do not like. However, the axiomatic theory is meant to judge the relative merits of index formulas when aggregating over heterogeneous commodities. When it comes to aggregating a homogeneous commodity over outlets, we think that the dimensionality axiom is irrelevant. Moreover, deflating the value index by the unit value index leads to the ratio of the quantities of the commodities sold as the implicit quantity index, and this is precisely what it should be in our view. And why worry about the fact that the unit value index does not pass the identity test?

A number of points should be kept in mind, however. The unit value index seems perfectly acceptable as a price index provided that the commodity in question is exactly the same across all outlets. The problem of course is to define ‘homogeneity’. Outlets provide certain services with their sales. So it can be argued that even physically identical products are not identical items when sold in different outlets due to different services that accompany the transactions. An extreme position would be to say that all transactions differ qualitatively and that homogeneity never occurs. Alternatively, the service can be seen not as a characteristic of the good sold, but as a product of its own, supplied free of charge. Notice the similarity with advertising. Although in one way or another services (like advertising) costs will be passed on to consumers, it is not at all clear that ‘better services’ always lead to higher prices - the opposite may be true in many cases. If consumers easily switch between outlets in response to relative price changes, this could indicate that the services provided don’t matter very much.

If a commodity is deemed homogeneous across outlet strata as well, the appropriate aggregation level to compute a unit value index would be the item level. Nevertheless, stratifying the set of outlets may sometimes be helpful in lowering sampling variances if only a part of the population is observed. Whether the standard errors of the unit value indexes are higher or lower than those of the other elementary aggregates is an open question. In any case, the unit value indexes are inclined to show a more erratic pattern. This is because they only use data in outlets with strictly positive sales. Standard practice on the other hand is to collect prices (if possible) also in outlets where the commodity is not sold that month (which is understandable from a pragmatic point of view but not as evident as it may look at first glance: what exactly is the economic meaning of the price of a commodity that is not sold?). Moreover, if the ‘price’ is temporarily unobservable, a price will usually be imputed.

4. Average transaction prices versus isolated price quotations

For most commodities in the Netherlands’ CPI prices are collected at a single point in time each month. In fact it is implicitly assumed that prices remain unchanged during the entire month. This could be a rather rigorous assumption. Diewert (1995) notes that “It should be evident that a unit value for the commodity provides a more accurate summary of an average transaction price than an isolated price quotation”. The time period over which the unit value is calculated should ideally be “the longest period which is short enough so that individual variations of price within the period can be regarded as unimportant”. If we understand him correctly, Diewert states that when prices change rapidly, the time period should be made shorter and price indexes should be calculated more frequently. In a situation of moderate inflation, however, price indexes will not be compiled more frequently than once a month. Since prices do not change continuously, there will exist a finite number, say K, of subperiods k within month t in which the prices $p_{gb}^{t,k}$ stay the same. We can write the monthly unit value as

$$(12) \qquad \frac{\sum_{k=1}^{K} v_{gb}^{t,k}}{\sum_{k=1}^{K} x_{gb}^{t,k}} = \sum_{k=1}^{K} z_{gb}^{t,k} p_{gb}^{t,k},$$

where $z_{gb}^{t,k} = x_{gb}^{t,k} \Big/ \sum_{k'=1}^{K} x_{gb}^{t,k'}$.

The unit value taken over the entire month weights the prices in the subperiods by the relative number of commodities sold. It seems to us that, in case of a homogeneous commodity, this provides an appropriate average transaction price; it reflects continuing price measurement in a very sensible way. Note that this presupposes homogeneity over time. A physically identical commodity compared at two moments in time does not necessarily have to be identical in an economic sense (it does not have to yield the same utility to consumers). Think for example of a Christmas tree bought a couple of weeks before Christmas or on Christmas Eve; surely these are different goods. There may even be a ‘time of the day effect’. Nevertheless, in practice such items have to be treated as identical - otherwise the problem of ‘seasonal’ commodities will become unmanageable. In the empirical analysis we assume that every specific coffee item sold in a certain outlet is homogeneous over the whole period under study. The base year price will be calculated as the unit value over that year. This differs from the method Statistics Netherlands uses, in which the base year price is the unweighted average of twelve monthly single price quotations.

We will compare coffee price indexes resulting from two different price concepts: the average transaction price, calculated as the unit value (the value in dfl per kilogramme) over the entire month, and the single price quotation. The latter will be estimated by the unit value over the week in which official price data collection actually took place, that is the week (starting with Sunday) in which the 15th falls. This should give an idea of the effect of the sampling in time.

5. Empirical results

The Nielsen data set contains weekly coffee sales over a period of 128 weeks, beginning with week 1 of 1994 and ending in week 24 of 1996, from 20 supermarkets in a Dutch urban area unknown to us. We have chosen to take a sample of specific coffee items that mimics the ‘basket’ in the official CPI. See De Haan and Opperdoes (1997) for details. Six (rather tightly described) items are selected, four of which are grinded coffees and two instant coffees; coffee beans are not in the sample. Two out of the four grinded coffees are so-called house brands. Official practice is to combine house brands and regard them as one (composite) item. The same goes for instant coffees. Looking at the average prices, however, our data suggest that the varieties within the composite items should be viewed as different items, and we will treat them as such.

Table 1. Laspeyres (higher level) price index numbers for coffee (1994=100)

Month	**Type of micro index**						
	Ratio of average prices	**Arithmetic mean (unwtd.)**	**Geometric mean (unwtd.)**	**Laspeyres**	**Paasche**	**Fisher**	**Unit value**
9501	117.7	117.7	117.7	117.6	117.5	117.6	116.8
9502	117.0	117.0	117.0	117.0	116.8	116.9	116.4
9503	118.5	118.6	118.5	118.6	118.3	118.5	119.3
9504	117.2	117.2	117.1	116.8	115.9	116.4	115.6
9505	120.4	120.4	120.4	120.2	120.1	120.2	119.8
9506	117.4	117.5	117.4	117.7	117.4	117.5	116.8
9507	120.1	120.2	120.1	120.2	120.2	120.2	119.8
9508	115.0	115.1	115.0	114.9	114.8	114.9	114.5
9509	114.6	114.7	114.6	114.9	114.6	114.7	114.2
9510	110.5	110.5	110.5	110.6	110.2	110.4	110.7
9511	102.9	102.9	102.9	103.1	102.9	103.0	102.8
9512	104.6	104.6	104.5	104.8	104.5	104.7	104.6
9601	99.6	99.7	99.7	99.6	99.5	99.5	99.5
9602	97.9	97.9	97.9	98.0	97.9	98.0	97.7
9603	100.1	100.1	100.1	100.2	99.7	99.9	100.3
9604	99.1	99.1	99.1	99.2	99.0	99.1	98.7
9605	99.4	99.4	99.4	99.5	99.4	99.4	98.9

9606	97.4	97.5	97.4	97.4	97.3	97.4	97.1

The 20 outlets are subdivided into two strata: 13 larger and 7 smaller supermarkets. For each of the 6*2=12 elementary aggregates we have calculated price indexes according to the seven different formulas mentioned before (including the unit value index); see table 1. Prices are calculated as unit values taken over the entire month, and base year prices are unit values over 1994. The micro indexes are direct (not chained) indexes. Remember that they are aggregated with fixed (1994) weights, which means that higher level substitution is always ruled out.

Perhaps with the unit value index as a minor exception, the various index formulas at the elementary aggregation level lead to almost identical results. This is due to the fact that there is very little variation in the prices of the specific items between outlets. Not only is the Netherlands' coffee market heavily dominated by one manufacturer and one supermarket-chain, but little price variation probably is a characteristic of coffee itself: prices seem to be determined largely by world market prices of coffee beans. The unit value index gives rise to a pattern which is slightly more erratic, although the long run trend very much resembles those of the other index series. Taking the Fisher formula as a reference point, the unit value index exaggerates price movements to a small extent.

Table 2 presents the differences between price index numbers using monthly chained micro indexes and those using direct micro indexes (from table 1). Only the (unweighted) arithmetic mean of price relatives and the Laspeyres, Paasche and Fisher indexes change by chaining, of course. Although the first is known to be prone to upward drift, the effect is rather small. Chaining does not seem to be a good thing to do, however, when using the Laspeyres or the Paasche formula. They have a large systematic upward and downward tendency, respectively. In contrast, the Fisher formula appears to be quite insensitive to chaining. A similar result was found by Reinsdorf (1995).

Table 2. Differences between Laspeyres (higher level) price index numbers for coffee (1994=100) with monthly chained micro indexes and direct micro indexes

Month	**Type of micro index**						
	Ratio of Average prices	**Arithmetic mean (unwtd.)**	**Geometric mean (unwtd.)**	**Laspeyres**	**Paasche**	**Fisher**	**Unit value**
9501	0.0	0.0	0.0	0.0	0.0	0.0	0.0
9502	0.0	0.0	0.0	0.1	-0.1	0.0	0.0
9503	0.0	0.1	0.0	0.2	-0.3	0.0	0.0
9504	0.0	0.2	0.0	0.7	-0.8	0.0	0.0
9505	0.0	0.4	0.0	1.7	-1.8	-0.1	0.0
9506	0.0	0.4	0.0	1.7	-1.8	-0.1	0.0
9507	0.0	0.4	0.0	2.0	-2.1	-0.1	0.0
9508	0.0	0.4	0.0	2.0	-2.0	-0.1	0.0
9509	0.0	0.4	0.0	2.1	-2.1	0.0	0.0
9510	0.0	0.5	0.0	2.5	-2.4	0.0	0.0
9511	0.0	0.5	0.0	2.5	-2.4	0.0	0.0
9512	0.0	0.9	0.0	2.8	-2.7	0.0	0.0
9601	0.0	0.7	0.0	2.8	-2.6	0.0	0.0

9602	0.0	0.7	0.0	2.8	-2.7	0.0	0.0
9603	0.0	0.7	0.0	2.9	-2.8	-0.1	0.0
9604	0.0	0.8	0.0	3.4	-3.2	0.0	0.0
9605	0.0	0.8	0.0	3.7	-3.4	-0.1	0.0
9606	0.0	0.7	0.0	3.6	-3.4	-0.1	0.0

Table 3 shows the differences between price index numbers where the price concept is the unit value over one week every month - thereby simulating current practices in taking isolated price quotations instead of average transaction prices - and those from table 1 where the price concept is the unit value over the entire month. In table 3A, base year prices are still measured as unit values over 1994. The differences are up to almost four index points in november 1995. Obviously, prices have been declining steadily during this month. As was mentioned earlier, it would be preferable to compute price indexes more frequently under such circumstances.

The figures in table 3B relate to indexes with base year prices calculated as the unweighted twelve month average of isolated price quotations, a method that is likely to be used when quantity information is lacking. This approach seems defensible only if demand is very inelastic. In general we may expect consumers to buy relatively less of an item at times when prices are high: they hoard, postpone their purchases or switch to cheaper brands. Particularly when (relative) prices change substantially during the base year, this method has the danger of overstating average base year transaction prices, so that price indexes will be understated. This is precisely what happened with coffee in 1994. The price indexes for coffee relating to table 3b are three to four points lower those relating to table 3a. However, month-to-month changes are (almost) similar using both methods of calculating base year prices.

Table 3. Differences between Laspeyres (higher level) price index numbers for coffee (1994=100) based on single price quotations and on average transaction

	A: Base year price is unit value over 1994			**B: Base year price is twelve month average**		
Month	**Type of micro index**					
	Ratio of average prices	**Arithmetic mean (unwtd.)**	**Geometric mean (unwtd.)**	**Ratio of average prices**	**Arithmetic mean (unwtd.)**	**Geometric mean (unwtd.)**
9501	0.0	0.0	0.0	-2.9	-3.8	-3.8
9502	-1.1	-1.1	-1.1	-3.9	-4.8	-4.8
9503	1.0	1.0	1.0	-2.1	-3.0	-2.9
9504	1.0	1.0	1.1	-2.0	-2.9	-2.8
9505	-0.4	-0.4	-0.4	-3.5	-4.3	-4.3
9506	0.5	0.5	0.5	-2.6	-3.4	-3.4
9507	-0.7	-0.7	-0.7	-3.8	-4.6	-4.6
9508	0.9	0.9	0.8	-2.2	-3.0	-3.0
9509	0.0	0.0	0.0	-3.0	-3.7	-3.7
9510	3.5	3.5	3.5	0.4	-0.3	-0.2
9511	3.8	3.8	3.8	0.7	0.3	0.3
9512	-0.5	-0.5	-0.5	-3.3	-4.0	-3.9
9601	-2.5	-2.5	-2.5	-5.0	-5.6	-5.5

9602	0.9	0.9	0.8	-1.8	-2.3	-2.3
9603	-0.2	-0.2	-0.2	-3.0	-3.5	-3.5
9604	0.7	0.7	0.7	-2.1	-2.6	-2.6
9605	0.7	0.7	0.7	-1.9	-2.6	-2.6
9606	0.9	0.9	0.8	-1.6	-2.3	-2.3

Table 4 shows the differences between price indexes that treat all coffee items, including of course the six selected ones, as identical goods and those from table 1. In the former case, we have one composite item, the price of which is calculated as the unit value (the value per kilogramme) of the store's sales on this composite item during the entire month. Base year prices are again unit values over 1994. The distinction between the two strata has been retained and the stratum price indexes are weighted with 1994 expenditure weights. Although the trend in the index numbers for the composite item comes close to that of the indexes in table 1, the index level is generally lower. This must be caused by too much aggregation.

Table 4. Differences between Laspeyres (higher level) price index numbers for coffee (1994=100) based on a composite item and on specific items

Month	**Type of micro index**						
	Ratio of average prices	**Arithmetic mean (unwtd.)**	**Geometric mean (unwtd.)**	**Laspeyres**	**Paasche**	**Fisher**	**Unit value**
9501	-3.5	-3.5	-3.5	-3.5	-3.5	-3.5	-2.8
9502	-1.8	-1.8	-1.8	-1.9	-2.0	-1.9	-1.5
9503	-2.0	-2.1	-2.0	-2.0	-1.7	-1.9	-2.9
9504	-2.6	-2.6	-2.6	-2.4	-2.1	-2.2	-1.6
9505	-2.3	-2.3	-2.3	-1.8	-2.0	-1.9	-1.8
9506	-2.4	-2.4	-2.4	-2.5	-2.3	-2.4	-1.6
9507	-1.8	-1.8	-1.8	-1.5	-1.6	-1.5	-1.4
9508	-2.0	-2.0	-2.0	-2.1	-2.2	-2.1	-1.7
9509	-1.7	-1.7	-1.7	-1.6	-1.4	-1.5	-1.0
9510	-2.8	-2.8	-2.8	-3.0	-2.8	-2.9	-3.4
9511	0.3	0.2	0.2	-0.7	-0.8	-0.8	-0.5
9512	1.6	1.6	1.6	1.6	1.8	1.7	1.9
9601	-0.4	-0.4	-0.5	-1.0	-1.3	-1.2	-1.1
9602	3.2	3.2	3.1	2.5	2.4	2.4	2.9
9603	0.6	0.6	0.6	0.9	1.1	1.0	0.6
9604	0.9	0.9	0.8	0.3	0.3	0.3	0.6
9605	1.4	1.3	1.4	1.5	1.6	1.6	2.1
9606	-0.6	-0.7	-0.8	-1.6	-2.6	-2.1	-1.9

6. Conclusions

Nielsen data allow us to compute micro price indexes according to various kinds of formulas. We have the opportunity to investigate which types of micro indexes clearly give biased results, for example by comparing these with Fisher (ideal) indexes. Unfortunately, at least in The Netherlands, coffee seems to be a bad choice for empirical research in this area. There is just too little variation in prices and/or price movements between outlets to find any substantial differences using different micro index formules. At best we can say that chaining should not be used in case of the Laspeyres (or Paasche) formula.

What does matter, however, is the price concept used. Taking unit values over one week every month instead of unit values over the entire month as the price concept shows differences that exceed by far the differences due to alternative elementary aggregate index formulas. The base year price concept is important too. Choosing an unweighted average of monthly price quotations (simulated by the unit values over one week) overstates the average base year transaction price and thus understates price indexes.

Appendix. Two approximations to the Laspeyres micro price index

Here it is shown under which circumstances the Laspeyres micro price index is approximated by the ratio of average prices (4) or the arithmetic mean of price relatives (5). Let us start by introducing our notation for some finite population statistics:

$\bar{x}_{gi}^{0} = \sum_{b \in B_{gi}} x_{gb}^{0} \Big/ N_{gi}$ is the stratum mean of base period expenditures on item *g*;

$\bar{p}_{gi}^{0} = \sum_{b \in B_{gi}} p_{gb}^{s} \Big/ N_{gi}$ is the stratum mean of prices of item *g* in period *s* (*s*=0,*t*)

$S_{x}^{2} = \sum_{b \in B_{gi}} \left(x_{gb}^{0} - \bar{x}_{gi}^{0} \right)^{2} \Big/ N_{gi}$ is the stratum variance;

$S_{xp^{s}}^{2} = \sum_{b \in B_{gi}} \left(x_{gb}^{0} - \bar{x}_{gi}^{0} \right) \left(p_{gb}^{s} - \bar{p}_{gi}^{s} \right) \Big/ N_{gi}$ is the stratum covariance between x_{g}^{0} and p_{g}^{s};

$CV_{i}\left(x_{g}^{0} \right) = \sqrt{S_{x}^{2}} \Big/ \bar{x}_{gi}^{0}$ is the stratum coefficient of variation of x_{g}^{0};

$\rho_{i}\left(x_{g}^{0}, p_{g}^{s} \right) = S_{xp^{s}}^{2} \Big/ \sqrt{S_{x}^{2} S_{p^{s}}^{2}}$ is the stratum correlation coefficient between x_{g}^{0} and p_{g}^{s}.

It can be shown that

(13) $$P_{gi}^{t}(R) = \left[\frac{1 + CV_{i}\left(x_{g}^{0} \right) CV_{i}\left(p_{g}^{0} \right) \rho_{i}\left(x_{g}^{0}, p_{g}^{0} \right)}{1 + CV_{i}\left(x_{g}^{0} \right) CV_{i}\left(p_{g}^{t} \right) \rho_{i}\left(x_{g}^{0}, p_{g}^{t} \right)} \right] P_{gi}^{t}(L),$$

and, with $p_{gb}^{0} x_{gb}^{0} = v_{gb}^{0}$, $p_{gb}^{t} \big/ p_{gb}^{0} = \pi_{gb}^{t}$ and obvious notation,

(14) $$P_{gi}^{t}(A) = \left[\frac{1}{1 + CV_{i}\left(v_{g}^{0} \right) CV_{i}\left(\pi_{g}^{t} \right) \rho_{i}\left(v_{g}^{0}, \pi_{g}^{t} \right)} \right] P_{gi}^{t}(L).$$

$P_{gi}^{t}(A)$ is an inferior approximation to the Laspeyres micro price index if there is a strong linear relation between outlet-specific base year expenditures and price changes. This may be the case with items subject to occasional promotion: outlets that sold *g* at a much lower price in period 0 than in period *t* will generally have low base period turnover but large price relatives. The correlation coefficient is expected to be negative and the Laspeyres index will be overstated. The problem grows when the index is chained month by month and there is 'price bouncing' (price oscillation), meaning that outlets launch promotion sales at different times; see for example Szulc (1987). In the latter case prices may permute in such a manner that the outlets simply exchange prices with each other. According to the permutation test, proposed by Dalén (1991), micro indexes should be insensitive to this kind of permutation. $P_{gi}^{t}(R)$ passes the test but $P_{gi}^{t}(A)$ doesn't. However, we do not see any reason why micro indexes should pass the permutation test.

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