ABSTRACT - RÉSUMÉ

Sensitivity Analyses for Harmonising European Consumer Price Indices

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As an integral part of the EUROSTAT work on the harmonisation of Consumer Price Indices a number of sensitivity analyses have been done for assessing comparability problems due to different methods used in different European countries.

So far sensitivity analyses have been done on the following micro-aggregation aspects:

- Formula choice
- Number of elementary aggregates
- Different ways of handling missing items
- Different ways of handling substitutions
- The age of item group weights

The methodology of and results from these analyses and other that will be done before the conference - will be reported.

Analyses de sensibilité pour l'harmonisation des indices des prix à la consommation européens

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Dans le cadre des travaux d'EUROSTAT, on a effectué un certain nombre d'analyses de sensibilité pour évaluer les problèmes de comparabilité que posent les différentes méthodes utilisées dans les différents pays européens.

Jusqu'à maintenant, les analyses ont porté sur les aspects suivants de la micro-agrégation :

- le choix de la formule
- le nombre d'agrégats élémentaires
- les différents moyens de traiter le problème des articles manquants
- les différents moyens de traiter les substitutions
- l'âge des poids des groupes d'articles

On présente la méthodologie de ces analyses et leurs résultats, et d'autres encore faites avant cette conférence.

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SENSITIVITY ANALYSES FOR HARMONISING EUROPEAN CONSUMER PRICE INDICES

The attached eleven studies each deal with one distinct problem in comparing different countries' CPI's with each other. Several of them, notably numbers 1,2,4,5,7 and 11 directly deal with micro-aggregation problems.

The purpose of all these studies is to assess the size of the disturbances resulting from different procedures in different European countries and the consequential need for harmonising these procedures.

Before the conference it is planned to do some further analyses, most importantly one on quality change. Also, a summary report synthesizing the findings in the separate analyses is planned for this year.

REASONS, WHY THE AVERAGE OF RELATIVES (R) IS UPWARDS BIASED

1) *Because of the unsymmetric nature of relative changes.* The absolute value of an upward change is larger than the absolute value of the corresponding downward change. For example going from 100 to 200 means + 100% but going in the opposite direction from 200 to 100 means a change of -50%. The average of these two changes becomes ½(100-50)=25% > 0 although nothing has happened!

The same thing could be explained by very simple algebra applied to price relatives. Let p be one price and c be the change in price. Then let the price in one outlet move from p to p+c and the price in another outlet move from p+c to p between the same two periods. The arithmetic "average" of the two relatives then becomes:

$$\frac{1}{2}\left(\frac{p+c}{p}+\frac{p}{p+c}\right)=1+\frac{c^2}{2p(p+c)}>1 \text{ as we have } p>0 \text{ as well as } p+c>0.$$

This is a simplified picture of what happens when outlets apply "bargain" pricing in a certain period and then return to its regular price.

2) *Because of the time reversal result.* If R_{01} denotes the forward change, R_{10} the corresponding backward change (from time 1 to time 0), p_{1k} and p_{0k} prices at times 1 and 0 respectively and n the number of observations, then we have

$$R_{01} \times R_{10} = (1/n)\Sigma_k\, p_{1k}/p_{0k} \times (1/n)\,\Sigma_k\, p_{0k}/p_{1k} \geq$$
$$(1/n^2)\{\Sigma_k\,(p_{1k}/p_{0k})^{1/2}(p_{0k}/p_{1k})^{1/2}\}^2 = (1/n^2)n^2 = 1$$

by the Cauchy-Schwarz inequality. Equality applies only if all price changes are equal.

This result could be seen as a generalisation of the simple one-observation result above to a situation with many observations. A return to the same price situation as in the base period via another price situation always leads to an index greater than 1 (100)!

3) *Because of the circularity test.* If we consider price ratios in single outlets to be stochastic and call them r_{01}, r_{12} and r_{02} for the respective time periods 0, 1 and 2, we then have $r_{02}=r_{01} \times r_{12}$. If we further have $R_{01}=E(r_{01})$, $R_{12}=E(r_{12})$, $R_{02}=E(r_{02})$, we obtain

$$R_{02}=E(r_{02})=E(r_{01}r_{12})=E(r_{01})E(r_{12})+Cov(r_{01},r_{12})=$$
$$R_{01}R_{12}+Cov(r_{01},r_{12})<R_{01}R_{12} \text{ if } Cov(r_{01},r_{12})<0$$

which is generally the case since the numerator of r_{01} is the denominator of r_{12}. (Otherwise prices would diverge indefinitely ...) This shows that the more frequently we chain, the larger will the R-type index be.

4) *Because the Laspeyres upward bias (compared to a cost-of-living or a Fisher index) is aggravated at the elementary level.* This is explained in detail by Moulton and Reinsdorf in recent BLS papers. If observations are self-weighted by base expenditure then R is a perfect representation of a Laspeyres index. But the Laspeyres index in itself is a much worse estimator of a COL at the elementary level than at the aggregated level (the consumer's propensity to substitute coffee in one outlet for coffee in another outlet is greater than his propensity to substitute petrol for coffee).

Laspeyres (L) and Paasche (P) indices without (with equal) weights would look like

$$I_L = \frac{\sum P_1 Q_0}{\sum P_0 Q_0} = \sum \frac{P_0 Q_0}{\sum P_0 Q_0} \frac{P_1}{P_0} \approx \frac{1}{n} \sum \frac{P_1}{P_0} = R_{01}$$

$$I_P = \frac{\sum P_1 Q_1}{\sum P_0 Q_1} = \frac{1}{\sum \frac{P_1 Q_1}{\sum P_1 Q_1} \frac{P_0}{P_1}} \approx \frac{1}{\frac{1}{n} \sum \frac{P_0}{P_1}} = \frac{1}{R_{10}}$$

We thus have an approximate Fisher index

$$I_F = \sqrt{I_L I_P} \approx \sqrt{\frac{\frac{1}{n}\sum \frac{P_1}{P_0}}{\frac{1}{n}\sum \frac{P_0}{P_1}}} = \sqrt{\frac{R_{01}}{R_{10}}} = F$$

which is possible to compute without weights. It can be shown by Taylor expansions that, just as Törnqvist and Fisher indices *with weights* approximate each other very well, F and G (the unweighted geometric mean) also approximate each other well.

It is also a simple matter to show that R≥F always holds so that *R always overestimates an approximate Fisher index.* This follows from the well-known inequality between arithmetic, geometric and harmonic means of positive real numbers.

CPI SENSITIVITY FOR FORMULA CHOICE

A number of sensitivity analyses have been done on Swedish CPI data as could be seen by the attached table. The purpose of the analyses was to gauge the sizes of differences between different formulae for elementary aggregates (EA). Four different formulae were tested:

A = *the ratio of average prices*,
R = *the average of price ratios*,
G = *the geometric mean of price ratios* and
RA = *the ratio of normed average prices* (used in Sweden's CPI)

EAs in two different price systems were used in the tests. DAPS (the DAily necessities Price System) is a system where both outlets and commodities are sampled by probability and covers goods normally found in supermarkets. LOPS (the LOcal Price System) is a system where outlets are sampled by probability but not commodities and it covers most other goods and some services typically bought in shops. In both these systems local interviewers are responsible for price collection. For both systems price changes from December year t-1 to December year t (the one-year link) were analysed. In the LOPS four different years were used (1990, 1991 and 1992 and 1993) but in the DAPS only 1992 and 1993.

The EA were divided into two groups, homogeneous and heterogeneous, according to the price variation within the EA. An EA was classified as HOMO if the price coefficient of variation was less than 50% and else as HETERO. The number of aggregates in each category is seen in the table (# AGG).

Four types of differences were analysed: A-G, R-G, R-A and RA-G. For each type of difference, type of EA and data set the MEAN and standard deviation (STD) of the differences were computed. The results are shown in the attached table. The interpretation is for example: In the LOPS 1990 there were 433 homogeneous EAs. Over the 433 EAs the (unweighted) mean of the index difference A-G was 0.48 and the standard deviation of the same difference was 2.23. The standard deviation is included in order to demonstrate that differences between e.g. A and G for single EAs could be substantial even though the mean over all EAs is rather close to zero.

The results support the following conclusions:

1) R consistently gives much larger results than the other formulae. The difference is often 2 index points or more.

2) On average A and G do not differ as much but it could well be 0.1-0.2 index points for a large aggregate. For single EAs and items - and particularly for heterogeneous item groups - the difference could, however, become dramatic.

3) The difference between RA and G is very small. The aggregate difference is well below 0.1 in all cases.

Table 1:

TYPE OF STATISTIC AND AGGREGATE	A-G	R-G	R-A	RA-G	# AGG
MEAN, HOMO, LOPS 90	0.48	2.05	1.56	-0.005	433
, LOPS 91	0.08	1.86	1.79	-0.010	329
, LOPS 92	-0.02	2.28	2.30	0.049	247
, LOPS 93	0.04	2.32	2.27	-0.068	258
, DAPS 92	0.13	0.66	0.54	0.019	1130
, DAPS 93	-0.09	0.94	1.03	-0.006	1126
MEAN, HETERO, LOPS 90	-1.63	3.22	4.85	-0.037	80
, LOPS 91	0.03	1.90	1.88	0.08	48
, LOPS 92	-0.99	2.57	3.56	0.15	31
, LOPS 93	0.12	2.22	2.10	0.14	29
MEAN, ALL, LOPS 90	0.15	2.13	1.98	-0.009	537
, LOPS 91	0.06	1.68	1.62	0.001	419
, LOPS 92	-0.12	2.21	2.33	0.057	291
, LOPS 93	0.05	2.21	2.16	-0.045	300
, DAPS 92	0.12	0.64	0.52	0.018	1167
, DAPS 93	-0.10	0.91	1.02	-0.007	1168
STD, HOMO, LOPS 90	2.23	2.62	2.72	0.34	433
, LOPS 91	1.55	2.43	2.94	0.32	329
, LOPS 92	1.31	3.48	3.71	0.32	247
, LOPS 93	2.05	2.66	3.44	0.31	258
, DAPS 92	0.61	0.94	1.04	0.08	1130
, DAPS 93	0.87	1.24	1.70	0.13	1126
STD, HETERO, LOPS 90	4.97	4.51	8.60	0.68	80
, LOPS 91	4.69	1.76	4.93	0.39	48
, LOPS 92	3.38	2.13	4.55	0.33	31
, LOPS 93	4.46	2.61	6.17	0.57	29
STD, ALL, LOPS 90	2.87	2.99	4.29	0.40	537
, LOPS 91	2.09	2.30	3.13	0.31	419
, LOPS 92	1.65	3.31	3.77	0.31	291
, LOPS 93	2.34	2.64	3.73	0.34	300
, DAPS 92	0.60	0.93	1.02	0.08	1167
, DAPS 93	0.97	1.25	1.81	0.14	1168

CPI SENSITIVITY FOR THE AGE OF ITEM GROUP WEIGHTS

Two types of sensitivity analyses based on Swedish CPI data have been done in order to investigate the effects of the "age" of weights.

1) One set of analyses was done based on a computer file with weights and monthly indices for about 300 item groups (codes) for the entire period 1981 to 1992.

The object of the analyses was to gauge the difference arising from the choice of weights. In order to do this we matched the December-December group indices with long-term weights for one up to five years earlier. We only used unchanged item groups (matching codes), so the comparisons are not based on all items. Certain matching codes do not have exactly the same content which disturbs the analysis to a certain extent. The results are shown in table 1 below. The presented figure is the difference between two weighted aggregate indices, based on matching codes, taking the one with older weights minus the one with the actual weights. As the time span between actual and old weights increases the number of matching codes decreases which, in the five-year case, decreases the aggregate weight for which the comparison is based down to about 75%.

In technical terms, the difference $D_{t,j}$ for the year t index compared with the index with weights that are j years old is:

$$D_{t,j} = \sum_i \left(w_i^{t-j} - w_i^t\right) \frac{p_i^{dec.t}}{p_i^{dec.t-1}}, \text{ where}$$

$$w_i^{t-j} = \frac{p_i^{dec.t-j-1} q_i^{t-j}}{\sum_i p_i^{dec.t-j-1} q_i^{t-j}}$$

Here, for item group i, $p_i^{m.t}$ and q_i^t denote prices in month m, year t and quantities year t, respectively. Note that here all weights are long-term with q representing the year after the price reference month.

Table 1: $D_{t,j}$ for different years t and weight lags j

t=index year	j=weight lag 1	2	3	4
1982	-0.02	--	--	--
1983	0.04	0.23	--	--
1984	0.11	0.02	-0.12	--
1985	0.06	0.08	0.02	-0.02
1986	0.06	-0.06	-0.15	-0.38
1987	0.39	0.36	0.52	0.56
1988	0.15	0.06	0.16	0.12
1989	0.03	-0.03	0.23	0.15
1990	0.17	0.24	0.18	0.74
1991	-0.07	-0.18	-0.33	-0.11

1992	-0.34	-0.33	-0.27	-0.32
1993	0.12	0.06	-0.19	0.11
MEAN	0.06	0.04	0.01	0.09

There is no clear pattern in these results. In certain instances the differences could be fairly large, in either direction. The negative differences for 1991 and 1992 are influenced by a rather radical tax and subsidy reform leading to increasing prices and weights for e.g. rent, telephone and other inelastic service prices. For other years there is a certain predominance of positive differences, particularly for the one-year comparison which is the most reliable one, indicating that older weights may lead to overestimated indices. This tendency is not necessarily significant, however.

2) Long-term indices (LTIX) were compared to short-term indices (STIX) for the same year in table 2 below. Both these indices are computed in the regular production process and they show the change from December year t-1 to December year t. In STIX, weights represent consumption values of the whole year t-1, while as in LTIX they represent year t. In both cases they are obtained from preliminary National Accounts figures divided into about 80 commodity groups. The difference between these indices is therefore mainly an effect of these different sets of weights although there are also some other factors coming into play. As expected newer weights generally result in a lower index number because of the substitution effect.

In technical terms, the difference D_t for the year t index is:

$$D_t = \sum_i \left(w_i^{STIX} - w_i^{LTIX}\right) \frac{p_i^{dec.t}}{p_i^{dec.t-1}}, \text{ where}$$

$$w_i^{STIX} = \frac{p_i^{dec.t-1} q_i^{t-1}}{\sum_i p_i^{dec.t-1} q_i^{t-1}} \text{ and } w_i^{LTIX} = \frac{p_i^{dec.t-1} q_i^{t}}{\sum_i p_i^{dec.t-1} q_i^{t}}$$

Note the distinction between these weights and those above. In STIX q represents the year before the price reference month.

Table 2: D_t for different years t

1979	1980	1981	1982	1983	1984	1985	1
0.14	**0.52**	**0.29**	**0.07**	**0.03**	**0.07**	**-0.06**	**0**
1987	1988	1989	1990	1991	1992	1993	M
0.25	**0.11**	**0.03**	**0.21**	**0.15**	**0.02**	**0.15**	**0**

♦

Both of these analyses are, in different ways, disturbed by other things than the pure weight effect. Taken together and adding the economic-theoretic reasons to expect higher index numbers with old weights, they do, however, support the view that the time span between weight renewals gives rise to important systematic differences in resulting index numbers.

CPI HARMONISATION REPORT 4:2
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CPI SENSITIVITY FOR NUMBER OF ELEMENTARY AGGREGATES

Two analyses based on Swedish data have been done in order to investigate the effects of the number of elementary aggregates.

1) Sensitivity analyses have been done on item group data for the entire period 1981 to 1992 (the same data as Report 3). The comparisons are between the official (OFF) index and an alternative index (ALT) based on somewhat more than half of the groups. The alternative index was constructed as follows:

> Item groups were organised into 2-digit levels and sorted by code. In this way similar items tend to come close together in the file. Then, within the 2-digit level, they were grouped two by two. Within each such group of two the item with the smallest weight was excluded from computations and the I_{ALT} was thus based on the remaining half of the items. If the number of items in the 2-digit level was odd, then the last and single one was retained. More exactly, I_{ALT} was based on about 55% of the items and 78 % of the weight.

This computation, shown in table 1 below, attempts to simulate what will happen if a country bases its index upon a smaller number of items by dropping the relatively smaller ones in weight.

Table 1:

Year	$I_{ALT} - I_{OFF}$	# GROUPS IN OFF	# GROUPS IN ALT	% OF WE
1981	-0.154	349	190	
1982	-0.308	350	192	
1983	-0.501	281	157	
1984	0.050	281	157	
1985	-0.118	284	157	
1986	-0.119	295	162	
1987	0.160	297	163	
1988	0.000	291	158	
1989	-0.309	288	158	
1990	0.329	291	160	
1991	0.166	281	155	
1992	0.605	303	164	
1993	-0.179	316	169	
MEAN	-0.030	301	165	

The results show no systematic difference but for single years the differences are sometimes substantial, from -0.5 to +0.6.

2) In the so called Local Price System prices for most items are measured in 2-3 industries (types of stores). Each industry is given a weight corresponding to its share of the total sales of the item group in question. An analysis was

done for 1992 and 1993 (December to December) by excluding all but the industry with the largest weight. The largest weight is often between 80 and 90%. Thereby two index numbers for the Local Price System were obtained, the difference of which are shown in table 2.

Table 2:

Year	$I_{ALT} - I_{OFF}$
1992	0.45
1993	-0.12

Given that it has only been possible to do these computations for two years, no hard conclusions can be drawn.

CPI SENSITIVITY MISSING FOR DIFFERENT WAYS OF HANDLING MISSING OBSERVATIONS AND SUBSTITUTIONS

Analyses have been done for the following two situations:

1. Missing items: An item in a certain outlet, for which there is a base price, is missing in month m and a substitute is not found.

2. Substitutes: For an item in a certain outlet a substitute is taken in month m.

Here we describe the results of these analyses.

Missing items

In the Swedish CPI the interviewers are instructed to first look for a substitute within the item definition when an item is not found in an outlet a certain month. If a substitute is not found we will have a missing observation for that month. The official procedure is then to leave that observation out of the compilations and compute average over the rest of the observations. This index called I_{OFF}. An alternative procedure is to carry forward the price of the previous month. The resulting index will be called I_{ALT}.

In the following table these two indices are presented for an index number based on a 12-month comparison (December t-1 to December t) for the Swedish local price system .

Table 1:

Year	I_{OFF}	I_{ALT}	Difference
1993	102.20	101.65	0.55
1992	97.89	97.64	0.25
1991	103.38	102.77	0.61
1990	105.52	104.55	0.97

Here we have an unequivocal and substantial systematic difference between the procedures. In years with high inflation (e.g. 1990), carrying forward the old price will lead to a significant underestimation of inflation.

Substitutes

When a certain item is discontinued and a substitute is selected, the interviewers have two options in the Swedish CPI.

1) Some item groups are subject to quality adjustment and the interviewer then assesses the quality difference in Skr between the new and the old item which is then added to (subtracted from) the base price.

2) In other item groups (mainly food) there is no quality adjustment. The price difference between the items will then, without reduction,enter the index computation.

Procedures 1 and 2 constitute the official index procedure and leads to the estimate I_{OFF}. Two alternative procedures have been tested:

A) We compare the two prices directly, without quality adjustment. We obtain the alternative index I_A.

B) We link the new price in without showing any price change. We then obtain the alternative index I_B.

In the following table the indices are compared.

Table 2:

Year	I_{OFF}	I_A	$I_A - I_{OFF}$	I_B	$I_B - I_{OFF}$
1993	102.20	102.17	-0.03	102.18	-0.02
1992	97.48	97.49	0.01	97.89	0.41
1991	103.03	103.09	0.06	103.06	0.03
1990	105.52	105.71	0.19	105.20	-0.32
1989	103.95	104.02	0.07	103.89	-0.06

These results do not provide unequivocal evidence for or against the existence of systematic differences between the analysed procedures.

CPI SENSITIVITY FOR INCLUSION/EXCLUSION OF CERTAIN ITEM GROUPS

Certain item groups in a CPI represent consumption that exists in all EU countries, yet they are at present included in some CPIs but not in others. How great problems for on comparability could this be expected to give?

A simple piece of algebra gives a first idea of the problem: The All Items Index could be written as the weighted sum of two indexes. One of these, I_S with weight w represents the potentially exluded item group and another I_R with weight 1-w represents all other items. We thus have:

$$\mathbf{I} = \mathbf{w}\,\mathbf{I}_S + (1-\mathbf{w})\mathbf{I}_R \text{ which gives } \mathbf{E} = \mathbf{I} - \mathbf{I}_R = \mathbf{w}(\mathbf{I}_S - \mathbf{I}_R) = \mathbf{w}\mathbf{D}, \text{ where}$$

D $= I_S - I_R$ is the difference between the item group index in question and the index for all other item groups and
E is the aggregate effect on the All Items index.

The critical limit for E is 0.1 In the following table we give some combinations of w and D giving E=0.1.

w	**D**
0.001	100
0.01	10
0.1	1

This means that an item group with a weight of, e.g., 0.01 will have to differ as much as 10 index points from the average of the index for it to influence comparability with the critical value 0.1. In the next table we show some real world effects for six item groups in the Swedish CPI for 1990-1993 which may be excluded in some other countries' CPIs. Three of these groups were included for the first time in 1993.

ANALYSIS OF THE EFFECT OF INCLUDING CERTAIN PARTICULAR ITEM GROUPS										
ALL ITEMS INDEX =		103.91			101.89			107.98		1
		1993			1992			1991		
ITEM GROUP	**1000W**	**D**	**E**	**1000W**	**D**	**E**	**1000W**	**D**	**E**	**1000W**
PACKAGED TOURS	6.8	-5.105	-0.035	8.5	5.204	0.044	8.2	-6.977	-0.057	8.7
AUTO INSURANCE	9.3	1.938	0.018							
HOME INSURANCE	4.5	1.597	0.007							
FUNERALS	2.6	-3.309	-0.009	2.6	4.371	0.011	2.4	5.794	0.014	2.4
LOTTERIES ETC.	14.7	-3.968	-0.058	12.1	-1.913	-0.023	11.4	-8.072	-0.092	11.4 -1
AIR TRANSPORT ABROAD	3.0	15.286	0.046							
ACCUMULATED EFFECT			-0.031			0.032			-0.135	

We see that the most influential item group is *lotteries etc.* Its influence - which in one case exceeds 0.1 - is due to its large weight and to the method used for it during these years which gave an index of 100 for all of the years 1990-1993. It is therefore unclear whether the accumulated effect should be seen as an effect of inclusion/exclusion or as an effect of the particular conceptual choices in the area of the lotteries item group. We see that none of the other effects (disregarding lotteries) reaches the critical level of ±0.1.

IS THE REPRESENTATIVE ITEM METHOD BIASED?

The most common method for selecting an item for pricing in a CPI is probably the representative item method. This method usually works in two steps. First, the central office works out a specification, then an interviewer selects a particular variety of this item in the shop for pricing. Different rules and guidelines exist for doing these selections. One could use a broad or a narrow specification and different rules of thumb could be used in trying to work out an item definition that will cover a large part of the item group as well as be possible to measure. For local selection, criteria could be, e.g., *most sold* or *lasts long* and in situations of substitution perhaps *most like*.

When discussing possible biases of this method, there must be a preferred method with which to compare it. This method is the method of probability sampling. In Sweden probability sampling for items is used in one subsystem (the Daily Necessities System) covering most food items and some other items sold in supermarkets. Here varieties are drawn from wholesalers' lists with probabilities proportional to old sales values for every variety (PPS sampling). The item sample thus drawn is very large, effectively covering some 800 items.

We have therefore studied the difference between an index number based on the whole probability sample and one based only on the item with the greatest sales value in the same item group. This analysis was done for two years, 1992 and 1993. For 1992 96 item groups with more than one item in the probalitiy sample were included and for 1993 115 item groups.

The following statistics were obtained for this difference over all item groups (a positive number means that the index based on all items was greatest) :

Year/Statistic	*1992*	*1993*
Mean	**-0.04**	**0.09**
Standard Deviation	4.81	8.38
Maximum	11.64	21.50
75% Quantile	2.41	3.99
Median	0.39	0.07
25% Quantile	-1.78	-3.44
Minimum	-19.67	-31.12
Number of item groups	96	115

The results show that, although the difference was sometimes very big in either direction for a certain item group, overall there is no systematic difference between the two methods. It must be stressed, however, that these results are valid only for food items in the Swedish CPI.

CPI SENSITIVITY FOR REGIONAL DEMARCATIONS

1. Commodities and services sold in outlets

In the Swedish CPI outlets are drawn randomly (with probability proportional to size sampling) from a business register representing the whole population of retail outlets. Thereby the whole country is represented and the outlets are in practice well distributed over the whole country.

The last columns of table 1 show the distribution of the outlet sample in 1993 in the local price system according to three criteria of regional divisions. The H regions denote different degrees of population density with H1 (Stockholm), H8 (Göteborg) and H9 (Malmö) being metropolitan areas. Aside of these areas H3 stands for densely-populated regions and H6 for sparsely-populated regions with H4 and H5 in between. R1-R8 constitute a division of the country into eight disjunct parts. Calculations for the Stockholm commune and for all metropolitan areas were added to simulate the situation where a country restricts price measurement to its capital or a few urban areas only.

Table 1 shows differences between indices based on outlets located in these regions. In the local price system (LOPS) these regional indices are generally based on a rather substantial number of outlets while for daily necessities (DANS) the number of outlets is sometimes very small; in two cases the index is based on one outlet only.

Table 1: Differences between regional indexes and the index for the whole country

Type of region	LOCAL PRICE SYSTEM 1990	1991	1992	1993	DAILY NECESSITIES 1993	Number of outlets in the 1993 sample LOPS	DANS
National index	106.3	102.9	98.0	102.3	101.7	548	74
H1	-0.0	-0.1	-0.5	0.2	1.1	112	19
H3	-0.1	-0.2	0.5	0.6	0.7	208	25
H4	1.1	1.5	0.1	-0.4	-0.9	95	17
H5	2.6	-0.4	2.0	-1.6	-1.0	36	2
H6	1.9	1.3	4.0	-2.8	0.8	23	1
H8	2.4	1.3	0.3	0.1	-1.2	45	6
H9	1.5	2.6	1.3	0.5	-2.0	29	4
R1	-0.1	0.1	-0.4	0.2	1.1	114	20
R2	-0.3	1.1	0.2	0.2	-1.0	95	4
R3	0.4	0.9	2.1	0.7	-0.5	55	8
R4	0.6	-0.1	1.1	-0.6	1.1	70	10
R5	1.2	0.1	0.1	1.4	-0.2	101	12
R6	2.9	1.1	0.3	4.3	-0.4	49	7
R7	1.2	1.2	3.8	-0.9	-0.5	23	1
R8	1.0	-0.5	1.7	1.2	-0.1	41	2
Stockholm commune	1.9	-0.9	-0.8	0.5	2.0		

All metropolitan areas	0.4	-0.2	0.5	0.1	0.3	176	28

It is not possible to discern any significant regional effects in table 1. In no case does the difference for all the four observations for a certain region go in the same direction from the national index. Our conclusion must therefore be that we cannot find any significant regional differences in the Swedish CPI data. From the data alone and with sampling variance in mind, it is even possible that the differences shown are entirely random.

This conclusion must, however, be qualified by (possibly special) Swedish circumstances. The outlets for which these computations are done are to a great extent united in chains of wholesalers and/or retailers with national price policies, even though there are still some price variations.

2. Rents

Another analysis was based on rents which are generally thought to be more variable by region than commodities found in outlets. This analysis is not based on CPI data but on an annual *dwelling and rent survey (DRS)*, conducted by Statistics Sweden. This survey uses a random sample of some 13000 rented apartments and is thus more reliable than the CPI renting survey (covering only some 1000 apartments) when it comes to regional breakdowns. (At the national level they move similarly.) More exactly the sampling error of the DRS' estimates is about 1% (95% confidence).

The DRS produces a regional breakdown into four categories:

Greater Stockholm (GS), roughly equivalent to region H1 above,
Greater Göteborg (GG), roughly equivalent to region H8 above,
Large municipalities (LM), consisting of municipalities with more than 75000 inhabitants and being outside the two metropolitan areas,
Small municipalities (SM), consisting of municipalities with less than 75000 inhabitants and being outside the two metropolitan areas.

In table 2a-c we present results for the period 1980-1993 based on this survey: The estimated variable is average rent for all apartments in a certain region. In contrast to the CPI, no quality adjustments are done.

Table 2a: Index numbers for rents in four regions of Sweden 1980-1993 (1980=100)

YEAR	GS	GG	LM	SM	SWEDEN
1993	369,9	380,3	384,8	363,9	371,4
1992	336,7	349,8	355,3	337,4	342,2
1991	307,5	322,5	326,3	308,9	313,8
1990	241,0	257,9	260,4	246,1	249,2
1989	208,1	224,3	231,0	216,0	218,1
1988	192,4	204,0	215,1	201,8	202,3
1987	178,5	189,0	200,3	189,7	188,7
1986	170,2	177,2	185,8	176,4	176,7
1985	161,5	166,9	172,0	164,0	165,2
1984	152,1	154,0	158,7	152,4	153,6
1983	143,0	143,8	149,4	143,3	144,4
1982	129,7	133,1	135,2	130,4	131,3
1981	111,7	112,9	115,5	114,7	113,8

In table 2a we have built up index series from 1980 for each region separately. It shows that regional differences have been moderate during the time period studied. In LM rents have increased only 4% more than in the national average and in GS 1½ % less. As for the average annual rate of increase it was between 10.6% in GS and 11.1 % in LM with 10.8% as the national average.

In table 2b we look at annual changes. Here we clearly see that regions tend to follow each other closely with rathe small variations in a certain year.

Table 2b: Annual changes from the year before.

YEAR	GS	GG	LM	SM	SWEDEN
1993	9,8	8,7	8,3	7,9	8,6
1992	9,5	8,5	8,9	9,2	9,1
1991	27,6	25,0	25,3	25,5	25,9
1990	15,8	15,0	12,7	13,9	14,2
1989	8,2	9,9	7,4	7,0	7,8
1988	7,8	7,9	7,4	6,4	7,2
1987	4,9	6,6	7,8	7,5	6,8
1986	5,4	6,2	8,0	7,6	6,9
1985	6,2	8,4	8,4	7,6	7,5
1984	6,4	7,1	6,2	6,3	6,4
1983	10,3	8,0	10,6	9,9	10,0
1982	16,1	18,0	17,1	13,7	15,4
1981	11,7	12,9	15,5	14,7	13,8

In table 2c we have focused on differences in annual change from the national average. Since the CPI weight of rented housing is more than 10%, we see (by multiplying the figures in table 2c by 0.1) that the potential effect from different choices on regional representativity, e.g. only including the capital of a country or only including metropolitan areas (GS+GM) in many years exceeds the critical limit of 0.1%.

Table 2c: Differences from the national average in table 2b.

YEAR	GS	GG	LM	SM
1993	1,3	0,2	-0,3	-0,7
1992	0,5	-0,6	-0,2	0,1
1991	1,7	-0,9	-0,6	-0,4
1990	1,5	0,8	-1,5	-0,3
1989	0,3	2,1	-0,4	-0,8
1988	0,6	0,7	0,2	-0,8
1987	-1,9	-0,2	1,0	0,7
1986	-1,6	-0,8	1,0	0,6
1985	-1,4	0,8	0,9	0,1
1984	0,0	0,7	-0,2	-0,1
1983	0,3	-2,0	0,6	0,0
1982	0,7	2,6	1,7	-1,7
1981	-2,1	-0,9	1,7	0,9

In Sweden the rental housing market was highly regulated during the time period considered here. In a country with a less regulated market, differences such as those presented above could be expected to be larger.

CPI SENSITIVITY FOR PRICE REBASING (and rounding) OF CONSUMPTION WEIGHTS

If, in the underlying definition of a CPI, the price and weight base periods are not the same a choice has to be made between:

$$I_1 = \sum \frac{w_b}{\sum w_b} \frac{p_1}{p_0} = \frac{\sum \left(q_b \frac{p_b}{p_0} \right) p_1}{\sum \left(q_b \frac{p_b}{p_0} \right) p_0} \qquad (1) \quad \textit{and}$$

$$I_2 = \frac{\sum q_b p_1}{\sum q_b p_0} \qquad (2),$$

where $w_b = q_b p_b$.

Both of these formulae differ from the Laspeyres' and other well-known index formulae. Bert Balk (1994) discusses the relation between them and their properties as approximations to a Laspeyres' index. However, Balk's conclusion that I_1 may be a better approximation to a Laspeyres' index than I_2 rests on false premises. His argument seems to be based on the assumption that, if we have three random variables X, Y and Z, positive correlations between X and Y and also X and Z imply a positive correlation between Y and Z. This is not generally true, however. It is true only if the former correlations are strong enough.

An advantage of I_2 is that it has a simple interpretation in basic index notation (written as a ratio of quantity weighted prices). I_1 implies that the quantity weights be recalculated based on the price change from b to 0. This price change may, however, have a connection with price change from 0 to 1 which is the target for measurement. An unhealthy tendency for weights to vary with the target variable is thereby introduced.

The following could, however, be said about the relation between (1) and (2): The weights in (1) will, in comparison with (2), decrease for products for which the prices increase more from b to 0. One could generally expect a tendency for products with great price increases in one period (b to 0) to increase less in the next period (0 to 1), i.e. a negative autocorrelation in a series of price changes. If so, small weights in (1) will be associated with small price changes from 0 to 1 and great weights with great price changes. Thereby a tendency for I_1 to become greater than I_2 will result.

Now to the empirical tests. It is unfortunately not possible (or at least very difficult) to compare (1) and (2) for many historical periods, since files with the original w_h are not saved for a long time in the Swedish CPI system . Only for the 1994 indexes the comparison could easily be done.

In the table below these calculations are presented. Note that b stands for the whole year of 1993, 0 for December 1993 and 1 for the current month. We see that the order of size of the difference seems to be greater than 0.1 with I_1 greater than I_2, as is to be expected theoretically.

MONTH	I2	I1	"Exact" I1	I1 - I2	"Exact" I1 - I1
9401	100.555	100.646	100.657	0.091	0.011
9402	100.976	101.121	101.116	0.145	-0.005
9403	101.429	101.551	101.542	0.122	-0.009
9404	101.934	102.064	102.052	0.130	-0.012
9405	102.108	102.246	102.235	0.138	-0.011
9406	102.100	102.235	102.232	0.135	-0.003

Two versions of I_1 were used: one based on the "exact" weights as given by the National Accounts and one based on initial rounding of these weights to whole ten-thousandths. This difference is small: in the order of 0.01. If a country uses cruder cruder weighting such as hundredths or thousandths the difference could get greater, however.

In later versions of this report more months will be included in the comparison.

Reference:

Balk, B.M. (1994): On the Choice of Weights for the Harmonized CPI. Statistics Netherlands, 19 January 1994.

CPI SENSITIVITY FOR DIFFERENT EDITING STRATEGIES

This report deals with the effects of different editing strategies, i.e. differences in the way national statistical agencies may check data and handle extreme observations. In the first section we will compare results based on raw and edited data. In the second section we will analyse effects of trimming, i.e. excluding extreme observations, since such procedures, more or less formalized are believed to exist in some countries.

1. Raw and edited data

In the Swedish CPI system data are checked and edited in two steps. In the first step, only some basic validity controls are done certifying e.g. that outlets and commodities have valid codes and reported prices are within very wide limits that are common for all commodities. These validity controls may be thought of as a minimum of editing, carried out in every country. In the second step more detailed checks are done; most importantly all price changes above a certain threshold (usually ±20% or ±40%) are listed for controls.

In order to establish the effect of the editing procedures, one wants to compare index computations from raw data with those from edited data. Since files with unedited data are not saved, it is not possible to make comparisons for historical periods. We will, however, do these comparisons for some months ahead. So far, computations are done for one month only - July 1994 (compared with December 1993). This gave the following results (*before editing* means after step one above but before step two and *after editing* means after step two):

Table 1: Effects of editing, July 1994

	Index before editing	*Index after editing*	*Differenc[e]*
Local price system	101.591	101.761	-0.17
Daily necessities	102.597	102.585	0.012

Computations will be done for at least the remaining months of 1994 and will be treated in later versions of this report.

2. Trimming

Little is known about the extent to which different countries accept extreme price changes in their CPIs. The Swedish practice is to accept all changes that are found to be correct after a validity check. In the most extreme cases price ratios of as much as 7 (or as small as 1/7) have been accepted. It is possible that in some other countries such extreme changes may be dismissed as being "unrepresentative".

In the table below a number of estimates of the one year CPI change in the two major price measurement systems LOPS (Local Price System) and DANS (Daily Necessities System) are given. Since the type of elementary aggregate formula used is potentially important, we have experimented with the three most important ones: A= the ratio of averages, R=the average of price ratios and G= the geometric average of price ratios. In addition to the untrimmed estimate (*None*), two trimmed estimates are given. Here a cutoff point of 2 (1.5) means that all elementary price ratios greater than two or smaller than 1/2 (2/3) have been excluded.

We see that in most cases the effects of trimming are substantial. For the R formula trimming generally decreases the estimate which is to be expected, since this formula treats price increases and price decreases unsymmetrically. The other formulas give smaller trimming effects and in both directions, but they are often greater than the critical level 0.1.

Table 2: Indexes based on different types of trimming

Formula/	LOPS				DANS	
Cutoff point	1990	1991	1992	1993	1992	1993
A						
None	105.20	102.30	96.88	104.68	97.50	102.00
2	105.17	101.98	97.48	104.33	97.58	102.06
1.5	105.15	102.16	98.17	104.16	97.92	102.03
R						
None	107.18	103.92	99.21	106.84	98.02	103.02
2	106.45	103.10	99.06	105.69	98.06	102.92
1.5	105.86	102.79	99.06	104.85	98.28	102.64
G						
None	105.05	102.24	97.01	104.63	97.38	102.10
2	105.03	101.88	97.56	104.21	97.47	102.13
1.5	105.08	102.09	98.23	104.02	97.85	102.08

3. Conclusions

Trimming rules may well have a systematic effect on CPIs, especially in connection with the R formula. If non-symmetric rules are applied with respect to price increases/decreases, the risk becomes, of course, even greater. A rule on trimming would therefore be appropriate in the harmonization exercise. Preferably trimming (interpreted as eliminating great price changes irrespective of whether they are real) should be outlawed all together.

Other than that no conclusions can yet be drawn on the effect of different editing strategies.

CPI SENSITIVITY FOR INCLUSION/EXCLUSION OF OUTLET TYPES

In a typical CPI sample of outlets, whether done by probability or judiciously, not all places where goods and services are purchased are eligible for selection in the sample. For example, food items may be sold in supermarkets, department stores, specialised food stores for fish, bread etc., kiosks, petrol stations, tobacco shops, market stalls, "street kitchens" or even flower shops and video rentals. Of these categories only a few would be included in a CPI and in the case of Sweden only supermarkets and department stores. Of course these different types of outlets will tend to keep different price levels but still the price change may not vary systematically.

In the Swedish CPI, many items are priced in department stores (shops selling a large variety of goods and having high turnover) and in one or two other outlet types and it is therefore natural to build a sensitivity analysis on a comparison between department stores and other types of outlets.

This has been done in the following two ways in the Local Price System for one-year comparisons (December-December).

First we look at the weighted aggregate (within LOPS) difference with and without department stores.

	1990	*1991*	*1992*	*1993*
Index excl. department stores	106.66	103.19	98.11	102.22
Index incl. department stores	106.34	102.88	97.97	102.32
Difference	0.32	0.31	0.14	-0.10
Weight excl. department stores	14.0%	16.7%	15.6%	15.0%
Weight incl. department stores	17.7%	20.5%	19.2%	18.0%

Then we look at the difference, item by item, between indexes for department stores and indexes for other outlet types and take the unweighted mean of those differences.

	1990	*1991*	*1992*	*1993*
Mean	-1.0	3.6	2.7	-1.4
t-value for mean=0	-1.2	2.8*	3.7*	-1.5
No of comparisons	219	174	139	142

*= the mean is significantly different from 0 at 5%-level.

These results are not conclusive for determining the existence of a long-term systematic difference between indexes with or without department stores. From an economic point of view it is unlikely for a certain type of outlet to increase prices more during any extended period of time. Still this may be possible for a period of, say, a few years. The data presented here hint at the possibility of such "medium-term" biases being greater than 0.1.