Statistics Netherlands

Division of Macro-economic Statistics and Dissemination
Support and development department

P.O. Box 4000
2270 JM Voorburg
The Netherlands

Generalized Fisher Price Indexes and the Use of Scanner Data in the CPI

Jan de Haan

Paper presented at the Sixth Meeting of the
International Working Group on Price Indices
Canberra, Australia
2-6 April 2001

Remarks:
The views expressed in this paper are my own and do not necessarily reflect the policies of Statistics Netherlands. I thank Bert Balk, Jörgen Dalén, Leendert Hoven and Mick Silver for helpful comments on earlier drafts and Eddy Opperdoes for research assistance. Email: jhhn@cbs.nl.

BPA-number: 0536-01-S-MOO
Date: March 20, 2001

Generalized Fisher Price Indexes and the Use of Scanner Data in the CPI

Abstract: Statistics Netherlands intends to use scanner data provided by retailers in compiling the CPI. This has two important advantages. First, taking a sample of items to estimate the commodity group price index for a particular type of outlet becomes unnecessary. Second, the Laspeyres-type index formula currently applied can be replaced by an index formula that is better suited for handling dynamic aspects such as commodity substitution and the introduction of new goods. The present paper suggests the use of a so-called generalized Fisher price index, based on a set of goods that is variable through time. This index contains prices of new and disappearing goods that cannot be observed directly and that should therefore be imputed. The relation with quality adjustment procedures is addressed as well.

JEL classification: C43, D12, E31.

Keywords: Consumer price index, imputation, new goods, quality adjustment, scanner data, substitution.

1. Introduction

In the past, empirical research on the Consumer Price Index (CPI), as well as the actual CPI calculation, had to rely on survey data. During the last couple of years the increasing availability of bar-code scanning data provides the opportunity to exploit the entire set of commodities (goods for short) belonging to a commodity group. This has stimulated the discussion on the preferred treatment of new and disappearing goods, which is a highly important aspect in the compilation of a CPI. Statistics Netherlands intends to use scanner data provided by nation-wide retailers in the production process of its CPI. The basic idea is to use the data on all goods at the commodity group level instead of taking samples, and to switch over from the currently used Laspeyres formula to the Fisher formula.

This paper suggests the use of a Fisher price index that explicitly accounts for new and disappearing goods, which will be called a generalized Fisher price index. The aim is not so much to develop any new ideas about CPI construction but rather to place various old ideas in the context of the generalized Fisher price index and to find out what we can learn from these ideas when we have scanner data at our disposal for an entire commodity group. Owing to the work of Diewert (1976) it has become generally accepted that a superlative price index like Fisher's reasonably approximates a cost of living index. In a challenging contribution, Balk (2000a) addresses the new goods bias and the substitution bias of the Laspeyres-CPI with respect to a cost of living index. Unfortunately there is no international consensus among statistical agencies on the question whether the theory of the cost of living index should be the underlying conceptual framework for the CPI (Triplett, 1999). Statistics Netherlands has always been, and still remains, one of the advocates of using this framework. However, the present paper takes a more pragmatic point of view. Hence it may appeal to those agencies that feel somewhat uncomfortable with the cost of living index methodology.

The remainder of this paper is organised as follows. Section 2 describes current practices used to estimate Laspeyres commodity group price indexes, and reviews some well-known drawbacks. Section 3 outlines the general idea behind the generalized Fisher price index. This index is based on a variable set of goods. It contains prices of new and disappearing goods that cannot be observed directly, which should be imputed. Section 4 discusses the main features of the generalized Fisher price index. Section 5 compares Balk's approach with ours and shows that both methods are compatible. Section 6 makes a case for using chained indexes. Section 7 presents empirical evidence using supermarket scanner data. Section 8 demonstrates how quality-adjusted unit values can link a disappearing good to a newly introduced one when both goods are close substitutes. Section 9 concludes and points to possible future empirical work in this field.

2. The Laspeyres price index estimator

The population Laspeyres price index for some commodity group I describes how the cost of purchasing the fixed set I^0 of goods belonging to I in the base period 0 evolves over time. For the current (comparison) period t this index is expressed as

$$P^t_{I,L} = \frac{\sum_{i \in I^0} p^t_i q^0_i}{\sum_{i \in I^0} p^0_i q^0_i}, \tag{1}$$

where p_i^s is the price of good i in period s (s= 0,t) and q_i^0 the quantity sold in the base period. Note that p_i^s will generally be some kind of average. With i deemed homogeneous, the unit value taken over (a sub-set of) all outlets that sell i – that is, the total expenditure on i divided by the total quantity bought – is the relevant average transaction price concept. See also Balk (1998).

In order to estimate $P_{I,L}^t$ the statistical agency draws a sample of items $\hat{I}^0$ from I^0. The usual estimator becomes

$$\hat{P}_{I,L}^t = \frac{\sum_{i \in \hat{I}^0} p_i^t q_i^0}{\sum_{i \in \hat{I}^0} p_i^0 q_i^0} = \sum_{i \in \hat{I}^0} \hat{w}_i^0 (p_i^t / p_i^0), \tag{2}$$

where the weight $\hat{w}_i^0 = p_i^0 q_i^0 / \sum_{i \in \hat{I}^0} p_i^0 q_i^0$ denotes the expenditure share of i with respect to the sample. In practice the weights are often proxies. If the sample size is small, the sampling variance of $\hat{P}_{I,L}^t$ may turn out to be rather high.[1] Normally a sampled item cannot be called a homogeneous good: the item description leaves enough room for the price collectors to select different varieties of the item in different outlets. The unit value taken over outlets will then not be a meaningful price concept, and the unit value index should not be used as an indicator of the item price index p_i^t / p_i^0. Hence, another type of item price index is called for. Since quantity and/or expenditure data at such detailed levels of aggregation is generally lacking, statistical agencies calculate item price index numbers simply from price data only.[2]

One of the problems associated with using the Laspeyres index (1) is that, by holding the quantities fixed at base period levels, substitution that takes place when consumers adjust their consumption behaviour in reaction to relative price changes is not taken into account. The question then arises how statistical offices treat new and disappearing goods. Some agencies regularly update their samples at least partially to account for new goods, like the BLS through annual sample rotation. Most agencies, however, re-sample only at base year revisions. Although

[1] Depending on the specific sampling design used, the estimator may exhibit small-sample bias (De Haan et al., 1999). Under sampling proportional to base period expenditure, the weights are left out from the estimation formula; they are implicitly reflected by the inclusion probabilities.
[2] The U.S. Bureau of Labor Statistics (BLS), for example, uses the ratio of geometric means of the prices observed in (a sample of) outlets to estimate most item price indexes. Statistics Netherlands still uses the ratio of arithmetic means but is considering changing over to geometric means (see De Haan and Opperdoes, 2001).

this might not be good practice, it is consistent with the Laspeyres price index (1), which ignores new goods altogether. Suppose that a statistical agency decides to adhere strictly to the Laspeyres principle. If it wants to keep the sample size fixed over time, the agency is forced to act when a sampled good disappears from the market. In that case it selects another item that replaces the 'old' one, probably the most similar item. The newly sampled item does not have to be completely new; it may already have been sold during the base period. To make the replaced item and its successor comparable, so that a quality change does not affect the price index, a quality adjustment must be made.

Quality adjustment seems a suitable name in the context of a survey-based Laspeyres price index estimator. But what if the entire set of goods is taken into consideration instead of a sample of items? There may not be an obvious one-to-one relation between a disappearing good and a successor. It may even be the case that the number of goods belonging to commodity group I decreases over time. The important thing to recognize is that quality adjustment methods are essentially imputation methods. Let $I^{0(D)}$ denote the sub-set of I^0 that disappeared in period t. The prices p_i^t for all $i \in I^{0(D)}$ in the Laspeyres price index (1) are 'fictitious' in the sense of being unobservable directly since there are no (monetary) transactions involved. These prices must be imputed.[3]

3. The generalized Fisher price index

Price indexes are usually defined on sets of goods that are fixed over time. In real life most sets are not fixed at all: apart from existing goods that disappear from the market, new goods enter also. Before turning to variable sets, we take a look at the Paasche price index. Let I^t be the set of goods belonging to commodity group I in period t and q_i^t the quantity sold of good i. The Paasche price index reads

$$P_{I,P}^t = \frac{\sum_{i \in I^t} p_i^t q_i^t}{\sum_{i \in I^t} p_i^0 q_i^t}. \tag{3}$$

[3] In a critical review of the harmonized index of consumer prices (HICP) constructed by European statistical offices, Diewert (1999a, p. 21) also notes that imputations of this kind cannot be avoided but claims that HICP regulations prohibit the use of imputations for non-monetary transactions. It is a matter of terminology, though. Quality adjustments do form part of the HICP methodology. Apparently the term imputations has been avoided to ensure that quality adjustments are allowed.

The set of goods that are new in period t will be denoted by $I^{t(N)}$. Base period prices p_i^0 cannot be observed directly for all $i \in I^{t(N)}$. These prices are fictitious – just like p_i^t for $i \in I^{0(D)}$ in the Laspeyres price index (1) – and must be imputed. The Paasche index excludes goods that disappeared after the base period, whereas the Laspeyres index ignores new goods. Both formulas also neglect substitution effects. To handle such dynamic changes we should look for an alternative price index formula. Using a symmetry argument, the square root of the product of the Laspeyres and Paasche price indexes seems a sensible candidate. This leads to the *generalized Fisher price index*, which is thus defined on a variable set of goods:[4]

$$P_{I,F}^t = \left[\frac{\sum_{i \in I^0} p_i^t q_i^0}{\sum_{i \in I^0} p_i^0 q_i^0} \frac{\sum_{i \in I^t} p_i^t q_i^t}{\sum_{i \in I^t} p_i^0 q_i^t} \right]^{1/2}, \tag{4}$$

in which $p_i^t = \hat{p}_i^t$ for $i \in I^{0(D)}$ and $p_i^0 = \hat{p}_i^0$ for $i \in I^{t(N)}$ are imputed prices.

Let $I^{0t} = I^0 \cap I^t$ denote the set of goods common to period 0 and period t.[5] It is assumed that $I^{0t} \neq \emptyset$. The Fisher price index defined on the set I^{0t} is

$$P_{I^{0t},F}^t = \left[\frac{\sum_{i \in I^{0t}} p_i^t q_i^0}{\sum_{i \in I^{0t}} p_i^0 q_i^0} \frac{\sum_{i \in I^{0t}} p_i^t q_i^t}{\sum_{i \in I^{0t}} p_i^0 q_i^t} \right]^{1/2}. \tag{5}$$

Using (5), the generalized Fisher index (4) can be decomposed into three factors:

$$P_{I,F}^t = P_{I^{0t},F}^t \left[\frac{\sum_{i \in I^t} p_i^t q_i^t / \sum_{i \in I^{0t}} p_i^t q_i^t}{\sum_{i \in I^0} p_i^0 q_i^0 / \sum_{i \in I^{0t}} p_i^0 q_i^0} \right]^{1/2} \left[\frac{\sum_{i \in I^0} p_i^t q_i^0 / \sum_{i \in I^{0t}} p_i^t q_i^0}{\sum_{i \in I^t} p_i^0 q_i^t / \sum_{i \in I^{0t}} p_i^0 q_i^t} \right]^{1/2}. \tag{6}$$

The second factor of (6) re-scales $P_{I^{0t},F}^t$ for the fact that the expenditures of new and disappearing goods have not been taken into account. $P_{I^{0t},F}^t$ and the re-scaling

[4] The choice for the Fisher (ideal) price index is sometimes justified on other grounds, particularly with reference to the test or axiomatic approach to measuring aggregate consumer price change. Diewert (1992) showed that the Fisher price index satisfies 20 'reasonable' tests, which is more than its competitors satisfy. These tests are based on a fixed set of goods; it is not necessarily true that the same holds if similar tests were developed for a variable set. The symmetry argument, on the other hand, applies to the generalized as well as to the ordinary Fisher price index.

[5] This set corresponds to Dalén's (1998) intersection universe.

factor only contain variables that can – at least in principle – be observed directly, for instance through scanner data. The third factor, which contains imputed prices, is needed to handle new and disappearing goods in the correct way.

The third factor of (6) deserves special attention. It can be rewritten as

$$\left[\frac{1+\sum_{i\in I^{0(D)}}\hat{p}_i^t q_i^0 / \sum_{i\in I^{0t}} p_i^t q_i^0}{1+\sum_{i\in I^{t(N)}}\hat{p}_i^0 q_i^t / \sum_{i\in I^{0t}} p_i^0 q_i^t}\right]^{1/2}. \tag{7}$$

The imputed prices $\hat{p}_i^t$ in the numerator of (7) might be obtained by making use of a quality adjustment method. In particular a hedonic regression could be run on data pertaining to the goods sold in period *t*, *i.e.* on (a sample of) the set I^t, apart from practical problems that may arise due to the lack of data on the goods' price-determining characteristics. With respect to the goods belonging to the set $I^{t(N)}$ in the denominator of (7) it is worthwhile distinguishing between 'more or less' new goods having technical characteristics of already existing goods, and completely new goods representing a new technology. The imputed prices $\hat{p}_i^0$ of the first sub-set of new goods could again in principle be estimated using hedonics, this time performed on data from the base period set I^0.[6] For the second sub-set of new goods, the completely new ones, the conceptual and practical problems are much bigger. These problems will be touched upon in section 6.

4. More about the generalized Fisher price index

The second and third factor of decomposition (6) both contain price and quantity data of new as well as disappearing goods. To gain further insight it can be helpful to separate new from disappearing goods. Rewriting the second factor of (6) as

$$\left[\frac{1+\sum_{i\in I^{t(N)}} p_i^t q_i^t / \sum_{i\in I^{0t}} p_i^t q_i^t}{1+\sum_{i\in I^{0(D)}} p_i^0 q_i^0 / \sum_{i\in I^{0t}} p_i^0 q_i^0}\right]^{1/2}, \tag{8}$$

and subsequently rearranging the product of (8) and (7) yields

[6] Fisher price indexes combined with the use of hedonic regression belong to the class of so-called superlative hedonic price indexes; see *e.g.*, Ioannidis and Silver (1997).

$$P^t_{I,F} = P^t_{I^{0t},F}\mu^t_I\lambda^t_I \tag{9}$$

as an alternative decomposition, in which

$$\mu^t_I = \left[\frac{1+\sum_{i\in I^{t(N)}} p^t_i q^t_i / \sum_{i\in I^{0t}} p^t_i q^t_i}{1+\sum_{i\in I^{t(N)}} \hat{p}^0_i q^t_i / \sum_{i\in I^{0t}} p^0_i q^t_i}\right]^{1/2} \; ; \; \lambda^t_I = \left[\frac{1+\sum_{i\in I^{0(D)}} \hat{p}^t_i q^0_i / \sum_{i\in I^{0t}} p^t_i q^0_i}{1+\sum_{i\in I^{0(D)}} p^0_i q^0_i / \sum_{i\in I^{0t}} p^0_i q^0_i}\right]^{1/2} .$$

The factors μ^t_I and λ^t_I can be seen as the effects of new and disappearing goods, respectively, on the generalized Fisher price index $P^t_{I,F}$. If the set I^{0t} of ongoing goods is very large compared to the sets $I^{t(N)}$ and $I^{0(D)}$, then μ^t_I and λ^t_I will both exhibit values close to 1, and $P^t_{I^{0t},F}$ might be an acceptable proxy for $P^t_{I,F}$. But we cannot expect this to be the case a priori. Notice that $\mu^t_I < 1$ if and only if

$$\frac{\sum_{i\in I^{t(N)}} p^t_i q^t_i}{\sum_{i\in I^{0t}} p^t_i q^t_i} < \frac{\sum_{i\in I^{t(N)}} \hat{p}^0_i q^t_i}{\sum_{i\in I^{0t}} p^0_i q^t_i},$$

which is equivalent to requiring

$$P^t_{I^{t(N)},P} = \frac{\sum_{i\in I^{t(N)}} p^t_i q^t_i}{\sum_{i\in I^{t(N)}} \hat{p}^0_i q^t_i} < \frac{\sum_{i\in I^{0t}} p^t_i q^t_i}{\sum_{i\in I^{0t}} p^0_i q^t_i} = P^t_{I^{0t},P}, \tag{10}$$

provided that there is at least one new good. Both the left-hand side and the right-hand side of (10) are Paasche-type price indexes, defined on the sets of new and ongoing goods, respectively. New goods are bound to exhibit relatively high base period prices had they been sold during that period. Thus we expect (10) to hold.

Similarly, $\lambda^t_I > 1$ if and only if

$$P^t_{I^{0(D)},L} = \frac{\sum_{i\in I^{0(D)}} \hat{p}^t_i q^0_i}{\sum_{i\in I^{0(D)}} p^0_i q^0_i} > \frac{\sum_{i\in I^{0t}} p^t_i q^0_i}{\sum_{i\in I^{0t}} p^0_i q^0_i} = P^t_{I^{0t},L}, \tag{11}$$

provided that at least one good has disappeared. The left-hand and right-hand side of (11) are Laspeyres-type price indexes, defined on the sets of disappearing and ongoing goods, respectively. Why should inequality (11) hold? Suppose no excess demand exists, *i.e.*, there is no supply rationing, so that consumers can freely buy any goods they like in any quantities. Hence, the disappearance of goods implies that demand has shifted away from goods with obsolete characteristics towards similar goods (either already existing goods or 'more or less new' ones), or that demand simply fell to zero. This must mean that the (imputed) prices $\hat{p}_i^t$ of the obsolete goods are 'too high' relative to the prices of the other goods.[7]

The demand-oriented view taken above fails if consumers cannot freely choose between substitutes, particularly in times of rationed supply. Suppose the manufacturer and/or retailer of some durable good decides not to sell a specific model of a specific brand any more. At the same time a new model is introduced, which has some new features. The price is increased, but by more than what could have been expected from the quality improvement. If consumers – for instance due to brand loyalty – hesitate to switch to other brands, which may have similar models, they face a ('quality-adjusted') price increase. In such cases too, the prices $\hat{p}_i^t$ of disappearing goods will be relatively high and λ_I^t probably exceeds unity.[8]

Recapitulating, in general we expect to find $\mu_I^t < 1$ and $\lambda_I^t > 1$. By using the Fisher price index based on the matched set of goods I^{0t} only, it is uncertain whether the generalized Fisher index (4) will be overstated or understated. In the short run, and especially if the set I^{0t} is large, $P_{I^{0t},F}^t$ may turn out to approximate $P_{I,F}^t$ rather well. In the longer run, on the other hand, when the number of goods belonging to I often grows, $P_{I^{0t},F}^t$ may well overstate $P_{I,F}^t$.

The phenomenon of new/disappearing goods is strongly related to that of substitution; it is merely a specific type of substitution.[9] Substitution effects and the effects of new and disappearing goods are usually treated separately, though. Substitution among ongoing goods (*i.e.*, within the set I^{0t}) is probably what most people have in mind when speaking loosely of substitution. This may be called substitution in a narrow sense. The first factor of decomposition (9), the Fisher price index $P_{I^{0t},F}^t$, takes account of this. The second and third factor, μ_I^t and λ_I^t, capture all other forms of substitution. Suppose that the statistical agency aims at

---

[7] Durable goods experiencing rapid technological change, such as personal computers, might be an example.

[8] Motor cars seem to be a good example.

[9] Balk (2000a, p. 2) puts it this way: "In each period the (representative) consumer is confronted with a set of available commodities and a corresponding set of prices. In each period the consumer makes his choice. In a later period some commodities are bought that were not or could not be bought in an earlier period; some commodities are no more bought; and some commodities continue to be bought although perhaps in different quantities. Thus there is substitution among continuing commodities, between new and continuing commodities, between continuing and discontinued commodities, and between new and discontinued commodities."

estimating the Laspeyres price index (1). It is easy to check that $P^t_{I,L} = P^t_{I^{0t},L}\lambda^t_I$. The difference between the expectation of $\hat{P}^t_{I,L}$ and $P^t_{I,F} = P^t_{I^{0t},F}\mu^t_I\lambda^t_I$ measures the bias of the Laspeyres price index estimator with respect to the generalized Fisher price index.

A straightforward way to split the bias into three additive terms is:

$$E\hat{P}^t_{I,L} - P^t_{I,F} = \left[E\hat{P}^t_{I,L} - P^t_{I,L}\right] + \lambda^t_I\left[P^t_{I^{0t},L} - P^t_{I^{0t},F}\right] + P^t_{I,F}\left[\frac{1}{\mu^t_I} - 1\right]. \tag{12}$$

The second term on the right-hand side of decomposition (12) can now be referred to as substitution bias in a narrow sense. We expect it to be positive. With $\lambda^t_I \geq 1$ in general the difference $P^t_{I^{0t},L} - P^t_{I^{0t},F}$ yields a lower bound to this component that can easily be calculated from scanner data. The third term of (12) measures the new goods bias of the Laspeyres price index. Since generally $\mu^t_I \leq 1$, this term has an expected positive sign also. The first term represents ‘statistical bias’, which mainly depends on the imputations for disappearing goods made in practice. The sign of this term, albeit unknown a priori, will likely be positive if the statistical agency substantially overstates the imputed prices $\hat{p}^t_i$ or, less precisely expressed, when it ‘undervalues quality improvements’. Boskin *et al.* (1996) suspected this to be the case for the U.S. CPI. Notice that when there is no item sampling involved, as with scanner data, the first term becomes $P^t_{I^{0t},L}[\hat{\lambda}^t_I - \lambda^t_I]$, where $\hat{\lambda}^t_I$ is a ‘non-survey’ estimate of λ^t_I.

5. A comparison with Balk’s approach

Balk (2000a) addresses the CPI’s substitution and new goods bias from a cost of living index perspective. He starts by making two assumptions: *i*) the preference structure of the representative consumer exhibits homotheticity, and *ii*) the unit cost (expenditure) function is of the CES (constant elasticity of substitution) type. Assumption *ii*) states that for any pair of goods i, j $(i \neq j)$ the demand elasticity of substitution $\sigma = -d\ln(q^t_i / q^t_j) / d\ln(p^t_i / p^t_j)$ is the same $(\sigma \geq 0; \sigma \neq 0)$ and that σ is also time-invariant. Assumption *i*) says that the optimal expenditure shares are independent of the utility level. Furthermore, it is assumed that *iii*) the actual expenditure shares in both the base period and the comparison period are equal to the optimal shares. A feature of Balk’s (2000a) approach – which makes it quite interesting to compare it with ours – is that the set of goods considered is variable but overlapping through time. Note that, in his section 5, he assumes a two-level structure in the consumer's preferences, which consists of unchanging commodity groups (the upper level) and changing ranges of commodities within these groups (the lower level). Within each group the elasticity of substitution is assumed to be constant, but between groups it may differ.

It is shown that the cost of living index, or rather the subindex for a certain commodity group *I*, can be expressed in a number of ways as the product of a conventional price index, defined on the set of goods common to the base period and the comparison period, and a factor depending on the change of the range of goods. Recast in our notation, the cost of living subindex can be expressed as[10]

$$P^{t}_{I^{0t},(.)}\left[\frac{\sum_{i\in I^{t}} p_i^t q_i^t / \sum_{i\in I^{0t}} p_i^t q_i^t}{\sum_{i\in I^{0}} p_i^0 q_i^0 / \sum_{i\in I^{0t}} p_i^0 q_i^0}\right]^{1/1-\sigma}, \tag{13}$$

where $P^{t}_{I^{0t},(.)}$ denotes the price index defined on the intersection I^{0t} according to some conventional formula (.). Although in theory the index formula to be used depends on the value of σ, we can safely choose the Fisher formula because all relevant types of index numbers will be much alike in practice. Thus

$$P^{t}_{I^{0t},F}\left[\frac{\sum_{i\in I^{t}} p_i^t q_i^t / \sum_{i\in I^{0t}} p_i^t q_i^t}{\sum_{i\in I^{0}} p_i^0 q_i^0 / \sum_{i\in I^{0t}} p_i^0 q_i^0}\right]^{1/1-\sigma} \tag{14}$$

can be used to approximate the cost of living subindex.[11]

Comparing (14) with (6) suggests that, if the generalized Fisher price index (4) is meant to approximate a cost of living subindex, the third factor of (6) – that is, expression (7) – must be

$$\left[\frac{1+\sum_{i\in I^{0(D)}} \hat{p}_i^t q_i^0 / \sum_{i\in I^{0t}} p_i^t q_i^0}{1+\sum_{i\in I^{t(N)}} \hat{p}_i^0 q_i^t / \sum_{i\in I^{0t}} p_i^0 q_i^t}\right]^{1/2} = \left[\frac{\sum_{i\in I^{t}} p_i^t q_i^t / \sum_{i\in I^{0t}} p_i^t q_i^t}{\sum_{i\in I^{0}} p_i^0 q_i^0 / \sum_{i\in I^{0t}} p_i^0 q_i^0}\right]^{\frac{1+\sigma}{2(1-\sigma)}}. \tag{15}$$

[10] For reasons of simplicity, a subscript for the commodity group has not been added to σ.

[11] This result is similar to that of Aizcorbe et al. (2000), who used a matched-item Törnqvist index instead of a matched-item Fisher index.

Now suppose that some new goods have been introduced, but no existing goods disappeared. Expression (15) then reduces to

$$\left[1+\sum_{i\in I^{t(N)}}\hat{p}_i^0 q_i^t \Big/ \sum_{i\in I^{0t}} p_i^0 q_i^t\right]^{-1/2} = \left[1+\sum_{i\in I^{t(N)}} p_i^t q_i^t \Big/ \sum_{i\in I^{0t}} p_i^t q_i^t\right]^{\frac{1+\sigma}{2(1-\sigma)}}. \quad (16)$$

Since the left-hand side is smaller than 1, expression (16) can only hold for $\sigma > 1$. Balk (2000a) arrives at the same conclusion about the value of σ, albeit in a rather different way.[12] The following relation can be derived from (16):

$$\frac{P^t_{I^{t(N)},P}}{P^t_{I^{0t},P}} = \frac{\eta^t}{\left(1+\eta^t\right)^{\frac{\sigma+1}{\sigma-1}} - 1}, \quad (17)$$

where $\eta^t = \sum_{i\in I^{t(N)}} p_i^t q_i^t / \sum_{i\in I^{0t}} p_i^t q_i^t$ denotes the ratio of the period *t* expenditure on new goods and ongoing goods. Note that the period *t* expenditure share of the new goods is $w_N^t = \eta^t/(1+\eta^t)$. Equation (17) describes the ratio of the Paasche-type price index of the new goods and the Paasche price index of the ongoing goods. Because $\sigma > 1$ and $\eta^t > 0$, this relative price change must be smaller than 1, which is in agreement with our earlier intuitive finding (10). It is an increasing function of η^t for given σ and a decreasing function of σ for given η^t. As an illustration, Table 1 contains percentage relative Paasche-type price changes, that is $[(P^t_{I^{t(N)},P} / P^t_{I^{0t},P}) - 1]100\%$, evaluated at various values of σ and η^t.

Table 1. Percentage relative (Paasche-type) price change of new goods

	$\eta^t = 0.025$	$\eta^t = 0.05$	$\eta^t = 0.10$	$\eta^t = 0.20$	$\eta^t = 0.40$
	$w_N^t = 0.024$	$w_N^t = 0.048$	$w_N^t = 0.091$	$w_N^t = 0.167$	$w_N^t = 0.286$
$\sigma = 1.2$	-92.0	-93.0	-94.6	-96.9	-99.0
$\sigma = 1.5$	-81.0	-81.9	-83.6	-86.6	-90.9
$\sigma = 2.0$	-67.5	-68.3	-69.8	-72.5	-77.1
$\sigma = 5.0$	-33.7	-34.1	-34.9	-36.4	-39.1

For example, Table 1 says that a substitution elasticity of 1.5 and a current market share of 4.8% of new goods (*i.e.* $\eta^t = 0.05$) corresponds to an (imputed)

[12] Most empirical work points to a value between 0 and 1. This work has been done on commodity group data. As a way out, Balk suggests the assumption of a two-level structure in the consumer's preferences mentioned above. He argues (p. 13) that requiring within-group substitution elasticities to be larger than 1 is consistent with the inter-group substitution elasticity being smaller than 1.

Paasche-type price change of new goods which lies 81.9% below the Paasche price change of ongoing goods. Notice the sensitivity of the relative price change of new goods to changes in the value of σ. As can be seen from expression (17), $P^t_{I^{t(N)},P} \approx P^t_{I^{0t},P}$ for extremely large values of σ. In this case all goods belonging to commodity group *I* become almost identical from an economic point of view, and we would expect to find equal price trends for all goods.

Suppose next that some goods disappeared from the market, but no new goods were introduced. Expression (15) now reduces to

$$\left[1+\sum_{i\in I^{0(D)}} \hat{p}_i^t q_i^0 / \sum_{i\in I^{0t}} p_i^t q_i^0\right]^{1/2} = \left[1+\sum_{i\in I^{0(D)}} p_i^0 q_i^0 / \sum_{i\in I^{0t}} p_i^0 q_i^0\right]^{\frac{\sigma+1}{2(\sigma-1)}}. \tag{18}$$

From (18) it follows that

$$\frac{P^t_{I^{0(D)},L}}{P^t_{I^{0t},L}} = \frac{\left(1+\xi^0\right)^{\frac{\sigma+1}{\sigma-1}}-1}{\xi^0}, \tag{19}$$

where $\xi^0 = \sum_{i\in I^{0(D)}} p_i^0 q_i^0 / \sum_{i\in I^{0t}} p_i^0 q_i^0$ is a shorthand notation for the ratio of the base period expenditures on disappearing goods and ongoing goods. Since $\sigma > 1$ and $\xi^0 > 0$, $P^t_{I^{0(D)},L} / P^t_{I^{0t},L}$ must be larger than 1, which is in agreement with our earlier intuitive finding (11).

Although our approach does not focus on the theory of the cost of living index, it seems that Balk's (2000a) method does not conflict with ours. He also proposes some simple methods for estimating σ. So perhaps his approach enables us to approximate the (generalized) Fisher price index from observable variables only, for instance according to (14), without having to rely on imputed prices. The ongoing goods should be classified according to the economic criterion of equal within-group substitution elasticities. But how should new goods be dealt with? "Commodities which are new in period *t* should be allocated to certain groups to the best of our (intuitive) knowledge. Their position can be reconsidered as soon as information about period *t*+1 becomes available, since at that time they belong to the set of ongoing commodities. This can eventually lead to a recomputation of the cost of living index for period *t* relative to period *t*-1." (Balk, 2000a, pp. 14-15). Hence, he proposes to impute unobservable substitution elasticities for new goods instead of imputing unobservable base period prices.

6. Chained indexes

New goods should be incorporated into the CPI as soon as possible. The natural method for doing this is to use chained indexes. Suppose that the direct index (4) is replaced by the product of the month-to-month generalized Fisher price indexes $P_{I,F}^{\tau/\tau-1}$:

$$P_{I,cF}^{t} = \prod_{\tau=1}^{t} P_{I,F}^{\tau/\tau-1} . \tag{20}$$

Chaining causes *i*) the set of ongoing goods not to shrink too much during time and *ii*) the sets of new and disappearing goods not to grow too large. Point *ii*) has obvious advantages for estimating the imputed prices $\hat{p}$. For example, it makes little sense to estimate in period *t* a base period price for a new good when the base period lies in the far past.

The chained index $P_{I,cF}^{t}$ is path dependent, whether or not the set of goods changes through time. That is, the index number in period *t* does not only depend on the prices and quantities in the base period 0 and the comparison period *t*, but also on the prices and quantities of all time periods $\tau = 2,....,t-1$ in between. Path dependency is not so problematic if the CPI is primarily viewed as a short-term indicator. Empirical studies, based on fixed sets of goods, tend to show no large systematic differences between direct and chained Fisher price indexes anyway.

According to Balk (2000b), one might view any chained price index as an approximation to the line integral Divisia price index. He argues that the Divisia price index can be given a meaningful interpretation using micro-economic theory of consumer behaviour if the usual assumption of a static preference ordering is relaxed. This assumption becomes less realistic when the time interval to which the index relates grows. Balk concludes that the Divisia price and quantity indexes can conceptually be viewed as the "ultimate economic price and quantity indices".

The imputed prices $\hat{p}_i^{\tau-1}$ for completely new goods introduced in period τ are called (Hicksian) reservation prices. The reservation price is the fictitious price that would reduce the demand for the product to zero had it been available in the period prior to its introduction. Some economists argue that reservation prices can really be computed in practice (see *e.g.*, Hausman, 1997), while others criticise the concept altogether (Hill, 1999). There appear to be no statistical agencies that are planning to apply the concept of reservation prices to their CPIs. One could doubt whether it would make much difference in practice provided that completely new goods enter the index shortly after their introduction on the market; neglecting those goods in (20) will in general have a minor impact. Though reservation prices are typically high, the quantities sold during the introduction period will be small. Moreover, the introduction of completely new goods does not happen frequently, at least not compared to 'more or less' new goods, which are merely new varieties of existing products.

Although through chaining new goods will be incorporated into the CPI as quickly as possible, calculating the commodity group price index from data of the set of goods common to period τ and period $\tau - 1$ only is not a perfect solution, even without the introduction of completely new goods. An obvious imperfection is the loss of information incurred. For example, the prices and quantities of new goods in period τ, while observable, are omitted from the computation. A related issue is that exact matching and chaining does not necessarily yield an accurate approximation to the chained generalized Fisher price index due to neglecting the effects of new and disappearing goods, similarly to the analysis in section 5 for the direct index.

7. Some empirical evidence

Statistics Netherlands receives scanner data directly from two of the largest Dutch supermarket chains, which are going to be used (according to current plans) in the actual computation of the CPI, starting from May 2001. Here, a selection has been made from the provisional database, covering expenditure and quantity data on 9 commodity groups during week 26, 1999 up to and including week 49, 2000. The weekly data were aggregated into data pertaining to 19 four-week periods (instead of calendar months, which would be required for the CPI). The European Article Number (EAN) identifies scannable products. Different EAN codes will be treated as separate goods, notwithstanding that different codes may represent items that are identical from the consumer's perspective; section 8 addresses this problem.

There are some 80 outlets in the sample.[13] For each EAN code sold in two consecutive periods a period-to-period unit value index has been calculated over all outlets as the ratio of the period-to-period expenditure index and the quantity

[13] For ease of computation the sample of outlets is held constant over time; it is thus a panel. In the CPI-database, however, the sample is allowed to change slightly over time. CPI-calculations will be based on (matched) outlets that are in the sample in two consecutive months.

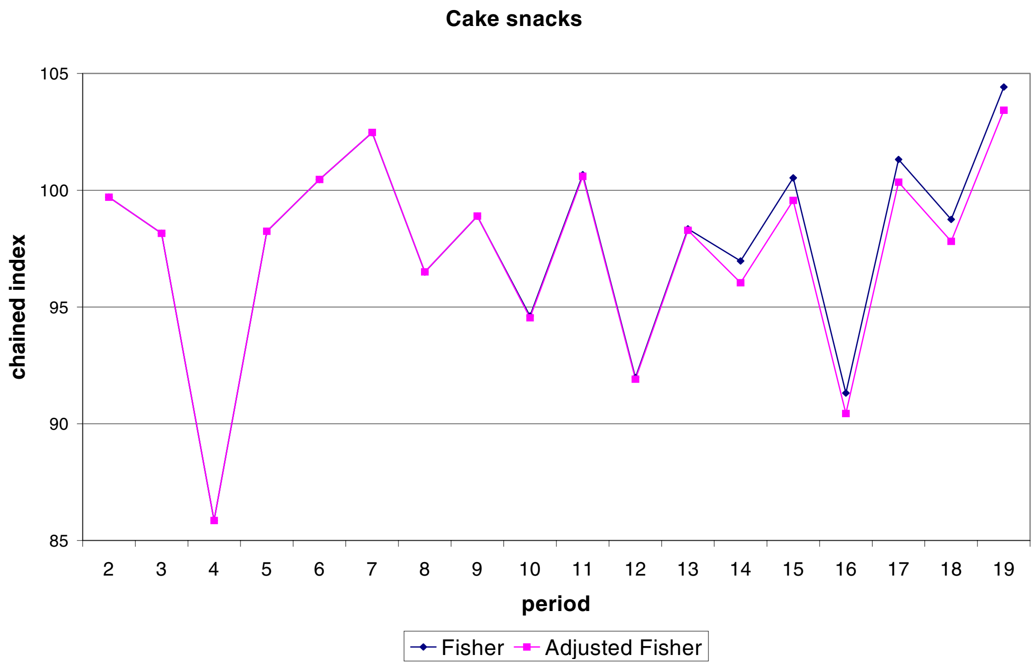
index (being the index of the number of scans). Next, commodity group period-to-period Laspeyres, Paasche and Fisher price indexes were computed using the unit value indexes of all EAN codes belonging to the group in question. No attempt was made to impute fictitious prices of new and disappearing EAN codes. That would have been an impossible task in case of fast moving consumer goods. What was done instead was to compute the period-to-period counterparts of expression (14), which approximates Balk's (2000a) CES-based price index and should thus be an approximation of a cost of living subindex. This index will be referred to as the adjusted Fisher price index. The within-group elasticities of substitution were estimated for each time period and according to three different methods following Balk's suggestions; see Opperdoes (2001) for details. Finally, chained indexes for the 9 commodity groups were calculated. Figure 1 presents the results; the chained Laspeyres and Paasche indexes are not shown.[14]

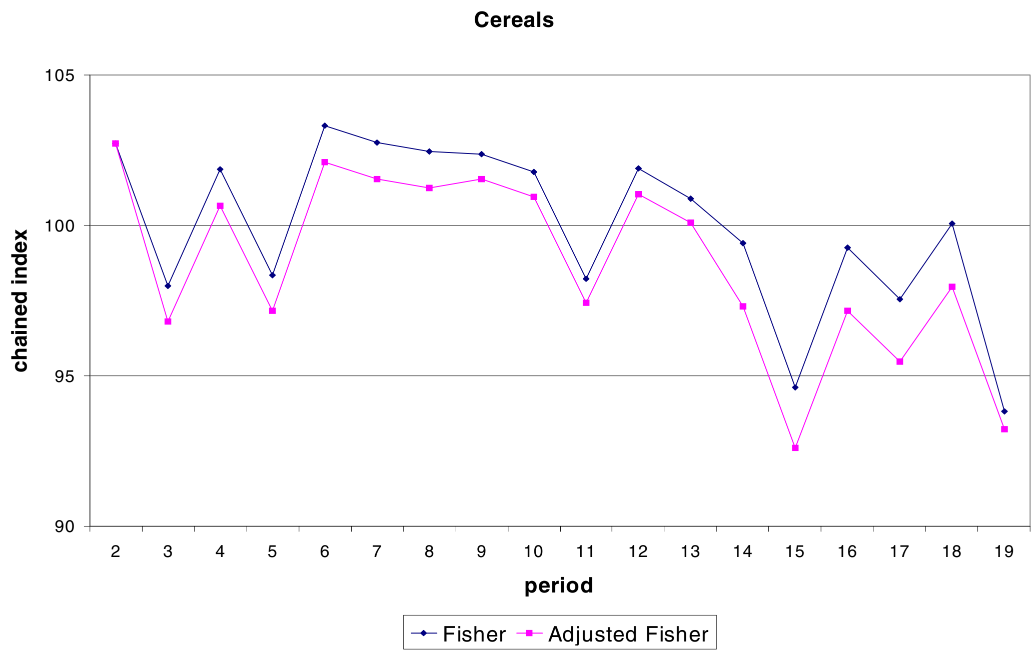
The difference between the Fisher price index and the adjusted Fisher price index differs among the 9 commodity groups. In 8 cases, the adjusted Fisher index lies below the unadjusted version. While for some groups the difference is rather small (particularly for cake snacks, cereals and crisps), for some other commodity groups it cannot be neglected. As a matter of fact, for tea, scents and yoghurt with additives the Fisher price index measures a price increase during the entire time interval, whereas the adjusted Fisher price index measures a price decrease.

Notice furthermore the volatility of the indexes. At first sight this erratic behaviour may come as a surprise, especially since there are such large amounts of data involved. Discounts, particularly those given to customer card holders, which are incorporated into the scanner turnover data are the probable cause. Consumers react strongly to discounts. Since discounts rest generally on popular products and because the use of customer cards is extensive (although varying across outlets), the impact of discounts on the commodity group price index numbers can indeed be quite large. Perhaps statistical agencies will be reluctant to accept such volatile price indexes. However, the erratic pattern reflects the use of average transaction prices at the level of individual goods coupled with the use of a superlative index formula to aggregate the goods' price indexes. Hence, it reflects actual consumer behaviour and describes a real phenomenon.

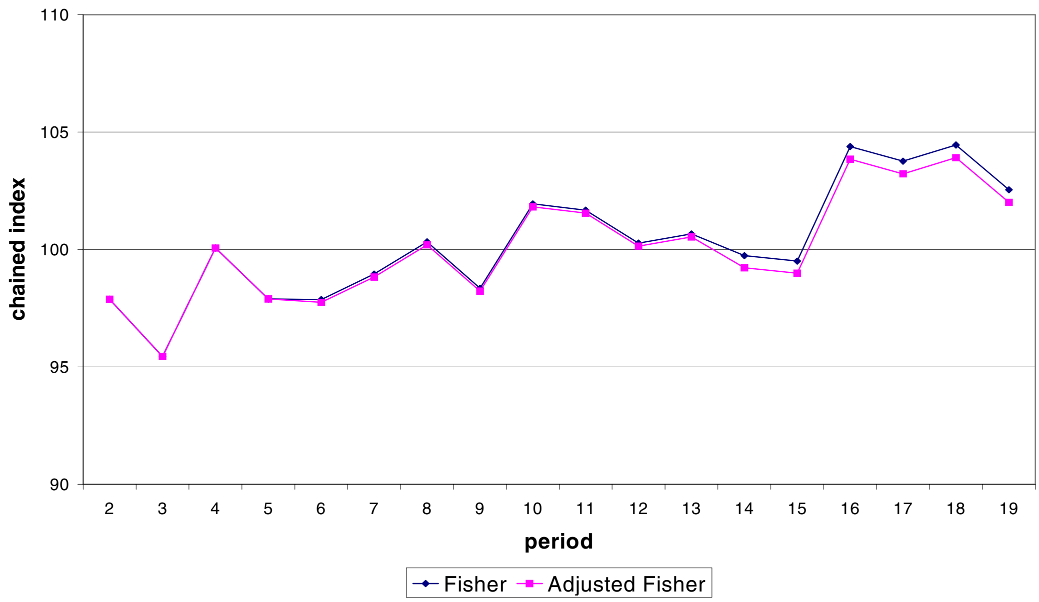
[14] The chained Laspeyres and Paasche price indexes exhibit strong upward and downward drift, respectively, which is a well-known result. The chained Laspeyres index numbers, for example, of baby's napkins and detergents are even well above 200 at the end of the time interval considered.

Figure 1. Chained price indexes for 9 commodity groups

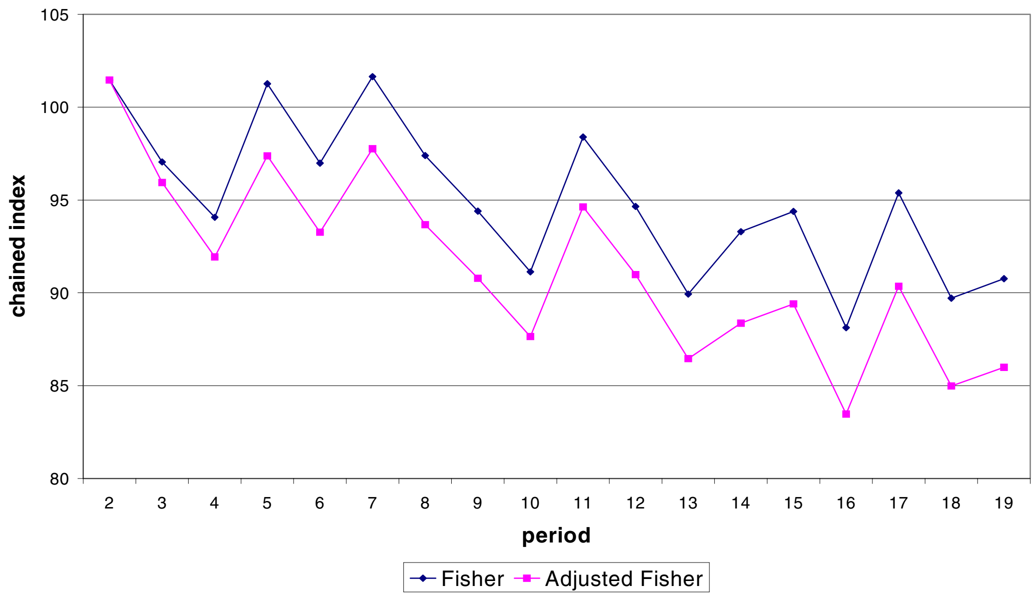




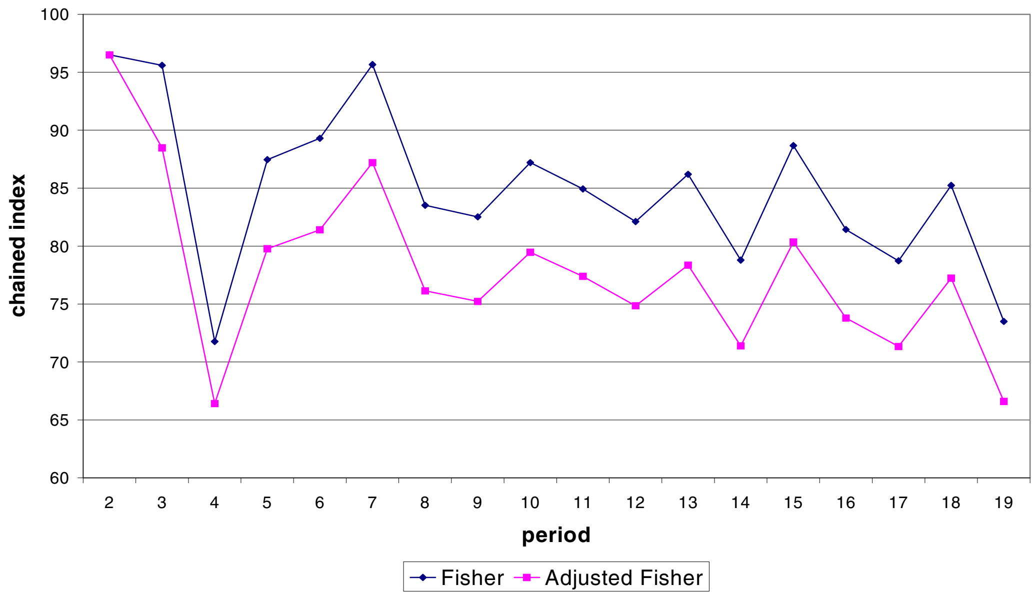
Crisps



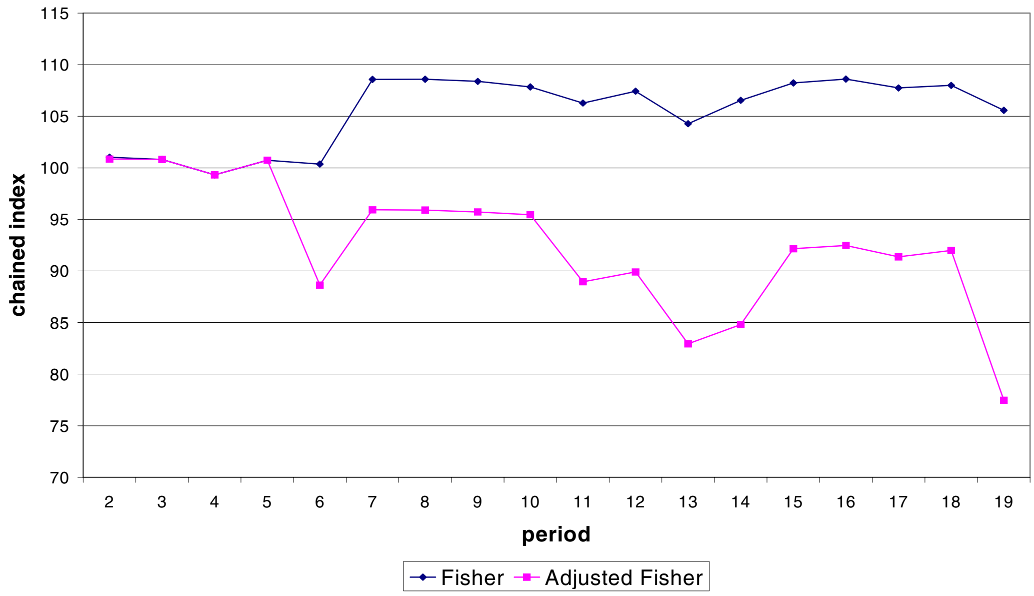
Detergents



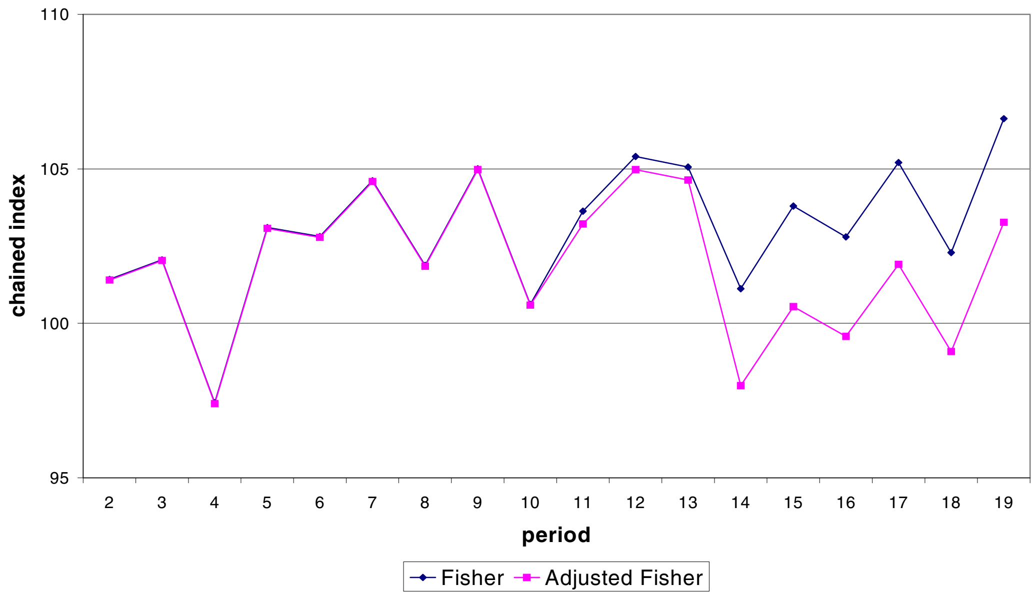
Disposable baby's napkins



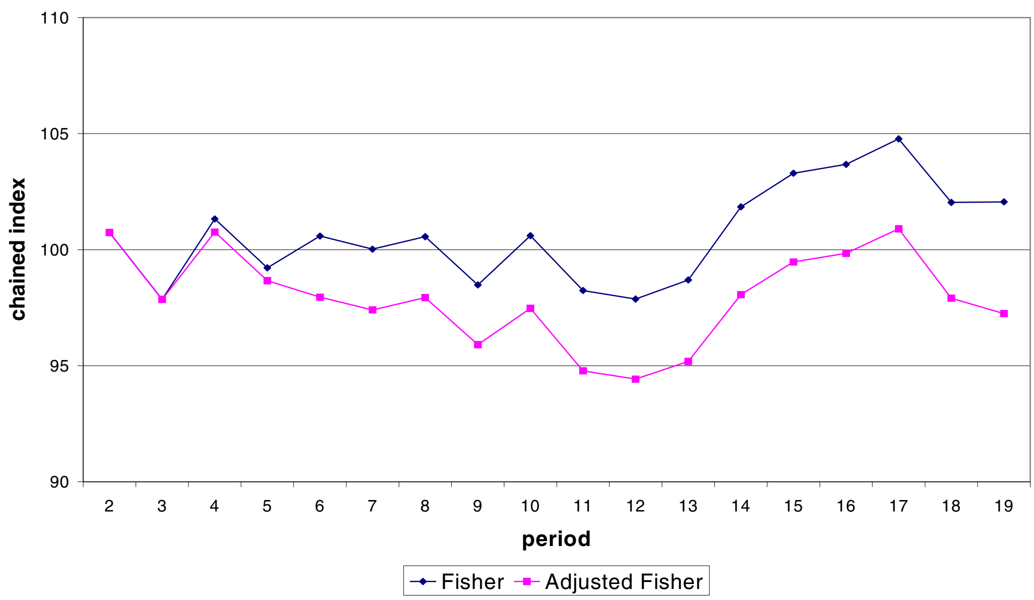
Scents



Soft drinks



Tea



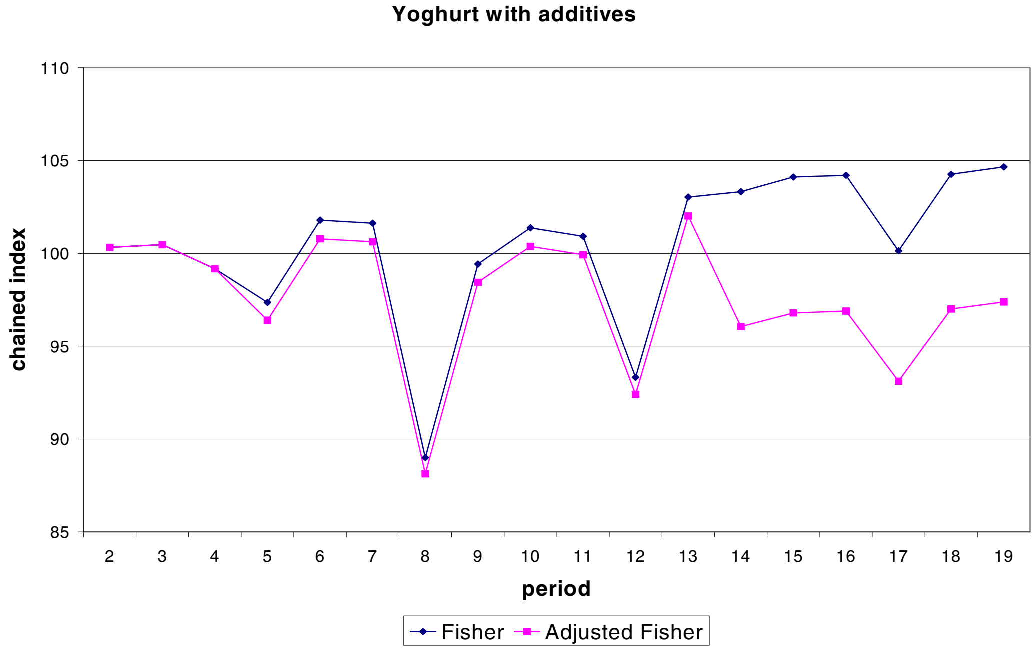


Figure 2. Aggregate chained price indexes

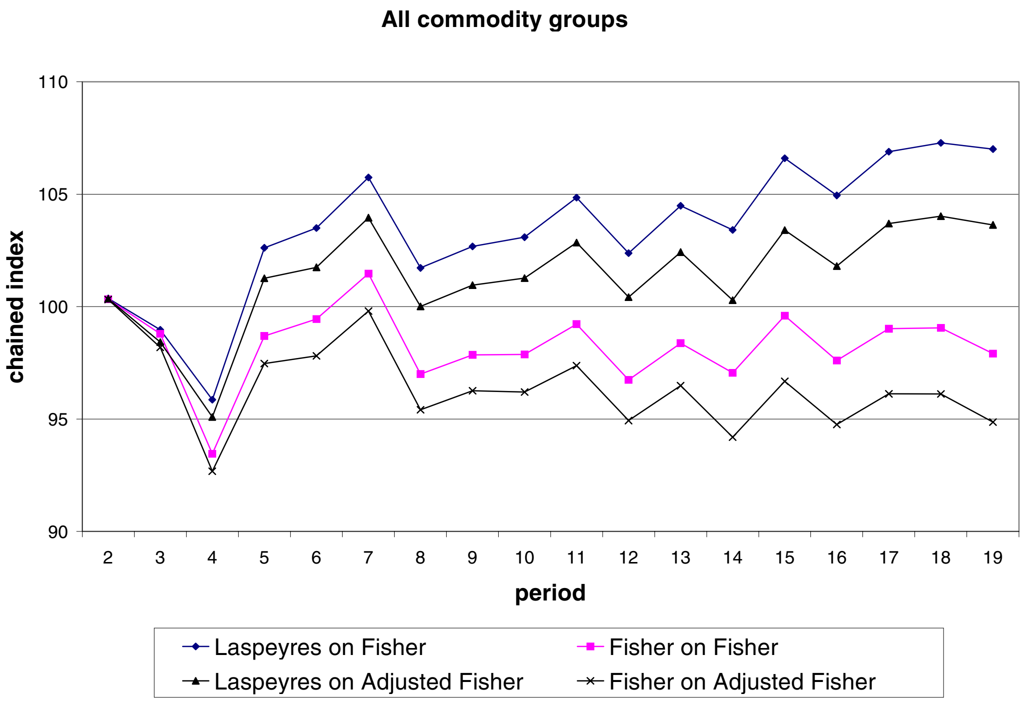


Figure 2 shows the aggregate price change of the 9 commodity groups. The period-to-period Fisher and adjusted Fisher commodity group indexes have been aggregated in two ways: using the Laspeyres formula (leaving expenditure weights fixed at period *t*-1 levels) as well as the Fisher formula. Time series were again obtained through chaining. As might be expected, the volatility of scanner data-based price index numbers diminishes at a higher level of commodity aggregation. Notice that the use of the chained Fisher formula to aggregate the period-to-period unadjusted Fisher commodity group price indexes points to a 2% price decrease between period 19 and period 1, whereas the use of the chained Laspeyres formula points to a 7% price increase. This difference illustrates the upward drift that is found using the latter formula, due to both the (upper level) Laspeyres formula as such and the chaining principle.

Table 2. Substitution elasticities and number of EAN codes

Commodity group	Substitution elasticities				Number of EAN codes			
	Mean*)	Stand. dev.	Min.	Max.	Mean	Stand. dev.	Min.	Max.
Cake snacks	4.5	1.2	2,6	6.6	73.5	2.7	70	79
Cereals	2.4	1.9	-5.4	4.0	18.4	0.7	17	19
Crisps	4.4	1.0	2.6	6.7	37.6	3.9	33	43
Detergents	5.0	1.4	2.3	8.4	50.3	3.2	45	56
Disposable baby's napkins	6.4	1.9	1.3	9.2	38.8	6.6	21	45
Scents	2.1	3.0	-3.5	7.3	22.4	1.9	20	26
Soft drinks	3.5	0.6	2.4	4.7	101.3	2.6	98	107
Tea	5.0	1.1	2.6	7.5	138.7	10.6	115	150
Yoghurt with additives	3.9	1.1	1.8	5.4	59.8	5.7	54	70

*) 10% trimmed mean

Let us now take a closer look at Balk's CES method. As was mentioned in section 5 he points to the fact that the goods should be classified according to the economic criterion of equal within-group substitution elasticities. That procedure has not been followed here. The individual EAN codes were classified according to a conventional classification scheme. Table 2 contains some statistics about the substitution elasticities and the number of EAN codes per commodity group. The 10% trimmed arithmetic mean of the estimated substitution elasticities has been used to compute the adjusted Fisher price indexes for each commodity group. In accordance with Balk's theory the trimmed means exceed unity, ranging from 2.1 (scents) to 6.4 (baby's napkins). However, for some commodity groups specific values below 1 and even negative values (cereals and scents) were found during certain periods. The large standard deviations suggest that the elasticities are not time-invariant at all. For all 9 commodity groups the number of products (EAN codes) varies during time and generally exhibit an upward trend.

The uncertainty about the 'true' value of the elasticities can be particularly a nuisance taking into account the sensitivity of the implicit ('fictitious') relative price change of new goods to changes in the value of the substitution elasticity shown in the example of Table 1. This raises another question. Is it realistic to assume, as Balk is doing, that a good's substitution elasticity be the same during its introduction and in later time periods? Moreover, it is difficult to understand how Balk's procedure of classifying goods should be carried out in practice. In any case, the procedure would probably bring about a lot of work for a statistical agency, and its practical feasibility can be doubted. Statistics Netherlands chose to implement chained Fisher indexes based on matched EAN codes only and not to use the adjusted CES-based version. Nevertheless, Figures 1 and 2 remind us that the use of the chained Fisher formula restricted to matched EAN codes can lead to biased figures.

8. Quality-adjusted unit values

Matching of EAN codes can create yet another problem, *i.e.* missing hidden price increases – or price reductions, but that seems less likely – and this is clearly undesirable. The EAN coding system may be too detailed so that different codes may well represent items that are identical from the consumer's perspective. In that case exact matching by EAN code would ignore any hidden price increase of a homogeneous item whose code has been changed. Although in section 3 it was argued that the fictitious, imputed prices of disappearing goods can (and should) be viewed in isolation from the fictitious, imputed prices of new goods, this might be a valid argument to link a disappearing good to a new one in a synthetic way – especially when these goods are very close substitutes and one could speak of a disappearing good and its natural successor. Thus, in order to reduce possible bias stemming from hidden price increases, the set of exactly matched EAN codes can be enlarged in a synthetic manner. This is the topic explored in this section.[15]

[15] The following part is a slightly changed version of section 3.4 in De Haan et al. (2000). Note that they also calculated chained Fisher commodity group price indexes (and various other types of price indexes as well, such as Törnqvist en Walsh indexes) based on exactly matched EAN codes from supermarket scanner data.

Suppose that good 1 has been sold in period $t-1$; it may be sold for some time during period *t* also, but we expect it to be sold no longer during period $t+1$. The quantities are thus $q_1^{t-1} > 0$; $q_1^t \geq 0$; and $q_1^{t+1} = 0$. Good 2, which is regarded as the obvious successor of good 1, was not yet sold in period $t-1$, but is sold in period *t* (and presumably in period $t+1$ as well). So we have $q_2^{t-1} = 0$; $q_2^t > 0$; and $q_2^{t+1} > 0$.

Suppose next that we can find a quality adjustment factor $\delta_{2/1}^t$ that serves to 'change the quantity bought of good 2 in period *t* into a quantity of good 1'. It is assumed that the representative consumer attains the same level of satisfaction from the consumption of one unit of good 2 as from the consumption of $\delta_{2/1}^t$ units of good 1. Hence, the consumer is indifferent between consuming q_2^t units of good 2 and consuming $\delta_{2/1}^t q_2^t$ units of good 1. An average price in period *t* of goods 1 and 2 can now simply be computed as

$$p_{2/1}^t = \frac{q_1^t p_1^t + \delta_{2/1}^t q_2^t \tilde{p}_2^t}{q_1^t + \delta_{2/1}^t q_2^t} = \frac{q_1^t p_1^t + q_2^t p_2^t}{q_1^t + \delta_{2/1}^t q_2^t}, \tag{21}$$

where $\tilde{p}_2^t = p_2^t / \delta_{2/1}^t$ might be called the quality-adjusted price of good 2; hence, $p_{2/1}^t$ might be referred to as a *quality-adjusted unit value*. To calculate the Fisher price index going from period $t-1$ to period *t* we should use the prices p_1^{t-1} and $p_{2/1}^t$, and the quantities q_1^{t-1} and $q_1^t + \delta_{2/1}^t q_2^t$. For the index going from *t* to $t+1$ we should use the prices $p_{2/1}^t$ and p_2^{t+1}, and the quantities $q_1^t + \delta_{2/1}^t q_2^t$ and q_2^{t+1}.

Expression (21) reduces to an ordinary unit value when $\delta_{2/1}^t = 1$. Such a situation can arise for example in case of a failed match, *i.e.* when an EAN-code changes while the goods in question are identical from the consumer's perspective – particularly when they are the same in physical terms. Notice that $p_{2/1}^t = p_1^t$ if $\tilde{p}_2^t = p_1^t$, that is when $\delta_{2/1}^t = p_2^t / p_1^t$, which implies the use of overlap pricing. If instead a hidden price increase occurs, so that $\tilde{p}_2^t > p_1^t$, then we have $p_{2/1}^t > p_1^t$.[16] Notice further that the average transaction price p_1^t is not defined in case $q_1^t = 0$. Taking $p_{2/1}^t = \tilde{p}_2^t$, which results from setting $q_1^t = 0$ in (21) had p_1^t been defined, is nevertheless the right solution.

---

[16] In a 'perfect world' the price difference between close substitutes would reflect the consumer's evaluation of the difference in quality when both goods are available at the same time, and overlap pricing would be the proper quality adjustment method. Consequently, a hidden price increase can only occur when the market is not 'in equilibrium' or when consumers are ill informed about the difference between the old and the new variety.

Determining the quality adjustment factors δ^t might be rather difficult in practice because of the limited product descriptions generally available in scanner data sets. There will likely be a need for collecting additional information on the goods' characteristics (compare footnote 16). For the sake of timeliness it could be useful to restrict such actions to disappearing goods with natural successors whose fractions of the commodity group's turnover in period $t-1$ exceed some threshold value – say 5 or 10%.

9. Summary and conclusions

CPI statisticians may sometimes get the feeling that they are faced with a paradox. On the one hand they should separate price from quantity changes, and to this end they try to keep many things fixed, for example the 'basket' of goods and services. On the other hand they should account for a changing consumption pattern of the representative household. The statisticians are aware of the discrepancy between a static Laspeyres-type CPI and the dynamic world they live in. To overcome this problem, pragmatic choices are being made. When a good of which the prices are collected disappears from the market, another good will be selected and a quality adjustment made to compute the desired 'pure' price change. When a completely new good appears it will usually be incorporated into the CPI somehow, albeit with a considerable time lag. To account for commodity substitution at a low level of aggregation, some agencies are using geometric means of price observations.

Since the CPI is a sample statistic, the statistical agency must have a view on the population statistic that the estimator represents. For instance, how should one interpret the Laspeyres price index estimator and the quality adjustments made when the number of goods in the population decreases? Looking at the population Laspeyres price index we can no longer speak of a one-to-one relation between a disappearing good and its natural successor. In this paper it was suggested to view quality adjustment methods as imputation procedures. Extending this idea to the Fisher price index, a 'generalized' Fisher price index was defined on a variable set of goods in which unobservable (fictitious) current period prices of disappearing goods and base period prices of new goods should be imputed. Note that, even if one is willing to accept the idea of the generalized Fisher price index as the ideal aggregator, there will always be room for controversy about the true value of the index number in a certain time period because different procedures to estimate the fictitious prices will lead to different outcomes.

The generalized Fisher price index rests on a simple symmetry argument; the Laspeyres price index uses forward imputation of disappearing goods' prices and the Paasche price index uses backward imputation of new goods' prices, and consequently their geometric mean uses both. It has been shown that the bias of the Laspeyres price index estimator – that is the difference between the expected value of the estimator and the true population value of the generalized Fisher price index – can be decomposed into three additive terms: statistical bias attributed to inadequate quality adjustment (*i.e.*, the use of incorrect fictitious current period prices of disappearing goods), substitution bias and new goods bias. Although the generalized Fisher price index was derived without explicit reference to the theory of the cost of living index, these three types of bias coincide with those which are usually distinguished when applying that theory.

A substantial part of this paper was devoted to the question under what conditions the generalized Fisher price index can be accurately approximated by the matched-item Fisher price index. The answer depends to a large extent on the expenditure size of the matched part, and possibly also on the nature of the goods in question and the prevailing market circumstances. Computing monthly-chained indexes instead of direct (bilateral) indexes has the advantage that the size of the matched part does not diminish during the course of time. In addition, monthly-chained price indexes can be viewed as approximations to Divisia price indexes, which are in some sense the 'ultimate economic price indexes'. Empirical work by Silver and Heravi (1999) on scanner data for washing machines seems to indicate that exact matching coupled with monthly-chained superlative price indexes leads to acceptable results. Still, a danger of obtaining biased results, *i.e.* of overstating or understating the generalized Fisher index, remains. This has been confirmed by our empirical analysis using an approximation of Balk's (2000a) CES-approach. A practical problem of using scanner data is that exact matching by the identifying EAN code disregards hidden price increases.

A well-known problem that has not been addressed in the paper, and for which a satisfactory solution does not seem to exist, is the presence of seasonal goods.[17] The CPI scanner database contains EAN codes, especially for fresh fruit,

[17] Diewert (1999b) recommends statistical agencies to construct three families of consumer price indexes to deal with this problem. The first index is defined over nonseasonal goods, the second index compares the prices of a certain calendar month with the prices of the same calendar month of the previous year, and the third index is an annual one (built up from the second index), which compares a moving total of 12 months with 12 base year months.

which disappear in certain periods and reappear in later periods. Chaining period-to-period indexes is probably not the right choice here, because reappearing EAN codes are treated as new goods.

Further research could be useful for helping statistical agencies to decide whether or not to use scanner data, particularly on durable consumer goods, in the production of their CPIs. One might, for example, attempt to follow the approach described in this paper using hedonic regression to estimate the fictitious prices – provided that the necessary data on the goods' price determining characteristics are available – and quantify the chained versions of decompositions (6) or (9). It would also be interesting to see how the results change if quality-adjusted unit values, estimated with hedonics, are used instead. Furthermore, one might try to calculate weekly instead of monthly-chained index numbers. By doing so, the size of the matched part will grow (in relative terms), but possibly at the expense of greater variability of the indexes. Research into sampling aspects can be relevant either. If a statistical agency decides to use scanner data, it may wish to do so on the basis of a sample of items. The question then arises how the sample should be drawn and what the statistical properties are of the chosen estimator.

References

Aizcorbe, A., C. Corrado and M. Doms (2000): "Constructing Price and Quantity Indexes for High Technology Goods", *Paper presented at the National Bureau of Economic Research Summer Institute 2000 session on Price, Output and Productivity Measurement*, Cambridge, July 31.

Balk, B.M. (1998): "On the Use of Unit Value Indices as Consumer Price Sub-indices", *Paper presented at the Fourth Meeting of the International Working Group on Price Indices*, Washington D.C., April 22-24.

Balk, B.M. (2000a): "On Curing the CPI's Substitution and New Goods Bias", *Research paper no. 0005* (Statistics Netherlands, Voorburg).

Balk, B.M. (2000b): "Divisia Price and Quantity Indices: 75 Years After, *Draft Research Paper* (Statistics Netherlands, Voorburg).

Boskin, M.J., E.R. Dulberger, R.J. Gordon, Z. Griliches and D.W. Jorgenson (1996): "Toward a More Accurate Measure of the Cost of Living", *Final Report to the Senate Finance Committee from the Advisory Commission To Study The Consumer Price Index.*

Dalén, J. (1998): "On the Statistical Objective of a Laspeyres' Price Index", *Paper presented at the Fourth Meeting of the International Working Group on Price Indices*, Washington D.C., April 22-24.

Diewert, W.E. (1976): "Exact and Superlative Index Numbers", *Journal of Econometrics*, 4, 115-145.

Diewert, W.E. (1992): "Fisher Ideal Output, Input and Productivity Indexes Revisited", *Journal of Productivity Analysis*, 3, 211-248.

Diewert, W.E. (1999a): "The Consumer Price Index and Index Number Purpose", *Paper presented at the Fifth Meeting of the International Working Group on Price Indices*, Reykjavik, August 25-27.

Diewert, W.E. (1999b): "Index Number Approaches to Seasonal Adjustment", *Macroeconomic Dynamics*, 3, 48-68.

Haan, J. de, E. Opperdoes and C.M. Schut (1999): "Item Selection in the Consumer Price Index: Cut-off Versus Probability Sampling", *Survey Methodology*, 25, 31-41.

Haan, J. de, P. van Poppel and C. Spaanderman (2000): "Using Scanner Data in the Compilation of the CPI: A Dutch Pilot Study", *Research paper* (Statistics Netherlands, Voorburg).

Haan, J. de, and E. Opperdoes (2001): "Meetkundig gemiddelde prijzen in de CPI" (Geometric Average Prices in the CPI), *Research paper* (Statistics Nether-lands, Voorburg).

Hausman, J. (1997): "Cellular Telephone, New Products and the CPI", *NBER working paper no. 598*2 (National Bureau of Economic Research, Cambridge).

Hill, P. (1999): "Inflation, the Cost of Living and the Domain of a Consumer Price Index", *Paper presented at the Joint ECE/ILO Meeting on Consumer Price Indices*, Geneva, 3-5 November.

Ioannidis, C., and M. Silver (1997): "Chained, Exact and Superlative Hedonic Price Changes", *Paper presented at the Third Meeting of the International Working Group on Price Indices*, Voorburg, April 16-18.

Opperdoes, E. (2001): "Using CES Functions to Improve the CPI: An Empirical Study", *Research paper* (Statistics Netherlands, Voorburg).

Silver, M. and S. Heravi (1999): "The Measurement of Quality-Adjusted Price Changes", *Paper presented at the Fifth Meeting of the International Working Group on Price Indices*, Reykjavik, August 25-27.

Triplett, J.E. (1999): “Should the Cost-of-Living Index Provide the Conceptual Framework for a Consumer Price Index?”, *Paper presented at the Fifth Meeting of the International Working Group on Price Indices*, Reykjavik, August 25-27.