

MEASUREMENT BIAS IN PRICE INDICES: AN APPLICATION TO THE UK'S RPI

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Abstract

This paper assesses the potential for systematic discrepancies between inflation as measured by consumer (or retail) price indices and one possible hypothetical ideal for monetary policymaking. It has long been suggested that measured price indices overstate the macroeconomist's view of inflation - as a result of both the weights used and the composition of the prices sampled. When monetary authorities choose to follow an explicit inflation target, the existence of a systematic upward bias between measured and 'true' inflation may affect the choice of measured target. In this paper, existing micro-level studies of bias in various components are reviewed and (where appropriate) used as the basis for extrapolation to the United Kingdom. By considering a range of assumptions, we derive a 'plausible range' of bias in inflation as measured by the UK's Retail Prices Index and its derivatives.

1. Introduction

"The Retail Prices Index (RPI) measures the change from month to month in the general level of prices charged to consumers across the range of the goods and services that they buy. As such, it is a measure of consumer price inflation. It does not measure the cost of maintaining a given standard of living - it is not a cost of living index."

National Audit Office, 'The Retail Prices Index', 1990, page 7.

UK policy-makers and others use the Retail Prices Index and its derivatives (such as RPIX and RPIY) for a variety of purposes. Table 1 (overleaf) lists various uses to which the Retail Prices Index is used alongside the form of price index ideal for that particular function. The ideal indices for each function are unlikely to be identical. Consequently, there is no reason to expect the RPI to correspond exactly with the hypothetical 'ideal' index for monetary policy and inflation targeting. This paper is concerned with systematic discrepancies between inflation as measured by derivatives of the Retail Prices Index and an ideal index for monetary policy purposes. The difference between the actual index and the ideal is known as a bias - clearly the bias is different for each use and does not necessarily imply any error in the construction of the published index.

Table 1
Uses of the Retail Prices Index

Retail price index used for:		Objective	Characteristics of ideal measure
Indexing:	Pensions	Maintain living standards	Constant cost-of-living measure Relevant to a particular subgroup of agents - pensioners
	Spending	Maintain spending levels	Constant cost-of-living measure Relevant to the spending patterns of one group of agents/ area of government
	Gilts	Offer a hedge versus inflation	
Cross country comparison		Macroeconomic analysis	A measure consistent across countries
Monetary policy		Maximise social welfare	<i>Either</i> <u>Empirical measure</u> An empirically determined measure of price change consistent with maximisation of long term GDP growth <i>Or</i> <u>Constant cost-of-living measure</u> Consistent with the theoretical stance that inflation in the cost of living is detrimental to growth.

Defining a price index for monetary policy

Monetary policy is generally concerned with the achievement of long run price stability. In the UK, this is manifested through the inflation target. A policy of inflation-targeting is founded on the premise that keeping inflation low and stable will maximise social welfare, which should equate with

maximisation of output over the long term. See Briault (1995) for a review of the theoretical and empirical literature on the costs of inflation.

Conceptually, the price index which is relevant for monetary policy might be described as a cost of living index defined as *a measure of the rate of change of prices for all goods and services currently being consumed*. As well as the acquisition cost, this should include the price of services which flow from durable goods consumption.

In practice, the price index chosen for the UK inflation target has been the Retail Prices Index excluding Mortgage Interest Payments (RPIX) which is not a perfect cost of living index:

- 1) it is a sample over a fixed selection of goods and services, from a fixed selection of stores,
- 2) the weights in the aggregate index are fixed,
- 3) it measures the asking price of goods and services, rather than those actually paid,
- 4) it is an acquisition cost index (i.e. it does not generally allow for the flow of services feature of durable goods purchase).

In its favour, an RPI-based measure is timely, consistent and has a high degree of public acceptability (as demonstrated by the wide variety of uses to which it is put). Furthermore, there appears to be no other index currently available which is closer to a cost of living measure, as defined above.

The discrepancies between the RPIX and a cost of living index mean that the RPIX is not an ideal index for monetary policy. So keeping inflation as

measured by the RPIX close to zero need not be consistent with maximisation of growth. There are two possible resolutions of this problem: either we estimate the optimal rate of RPIX inflation or we estimate the deviation of RPIX inflation from the ideal.

There have been some attempts to estimate the optimal level of inflation, i.e. that which is consistent with growth maximisation. However, while a negative correlation between inflation and growth has been detected in cross-country studies [see for example Barro (1995)], the results have not been sufficiently precise to define an optimal level of inflation.

The approach adopted in this paper is to evaluate a plausible range for bias in the Retail Prices Index compared with a cost of living index.¹ Since any bias has probably varied over time, our evaluation is based on conditions prevailing in the 1990s. Although we would like to cover all of the potential differences outlined above, in practice the measurement of the price of the flow of services from durables consumption does not appear to be tractable for the whole index. Consequently, we concentrate on the bias in the RPI as a measure of acquisition cost. The limited evidence from available micro studies suggests that allowing for the flow of services could substantially increase the upper end of our suggested range.

The various possible sources of bias between a cost of living index and the RIP are outlined in detail in Section 2 of this paper. In Sections 3 through 6 we assess each bias in turn. In each section, we assess the existing micro-level

¹ RPIX and RPI inflation rates are assumed to be biased to a similar extent.

literature and broader estimates of bias in the United States and Canada. This is followed by an estimate of a 'plausible range' for the UK. The ranges are derived from simple simulations and extrapolation from micro-level studies. The technique does not pretend a high degree of accuracy, but it does serve to expose the extreme assumptions necessary to reach a high level of bias. For each bias, the plausible range suggested covers a fairly broad range of assumptions.

There is no reason to presume that any difference between inflation as measured by the RPIX and under a hypothetical index has been constant over time. In addition to changes in the Retail Prices Index construction 'technology', any deviation may vary with the state of the economy. Given the vagaries of the discrepancy, we have estimated a 'mean' bias using inputs plausible for the 1980s and 1990s. In Section 7, we discuss the variance of the bias - at a qualitative level.

2. Sources of Bias

Systematic differences between consumer price indices and a conceptual cost of living index arise both because of the fixed nature of the sample on which price observations are taken and the fixed nature of the weights used to calculate an index from the observations. With continually changing preferences and technology of supply, the set of goods available varies over time. If the sample used to construct the price index does not shift in line with the available set, then the index can become distorted. The ensuing biases are termed 'compositional'. The weights used to calculate an index from the set of observations are fixed over some period. But, even if the composition is accurate, optimising agents can substitute in response to relative price movements, so that an 'ideal' set of weights will also respond to relative price movements. The ensuing discrepancy between measured and conceptual indices is termed 'substitution' bias.

Compositional Bias

'Compositional' bias derives from the failure of the RPI to deal with changes in the physical characteristics of goods consumed. There are two aspects - the ability to deal with entirely new goods, and the ability to allow for changes in the quality of existing goods. At a theoretical level, the two are quite similar - any change in quality might be regarded as the introduction of a new good at the expense of an old one. However, it is useful to distinguish in practice between the problems involved in introducing products with no historical comparison and the problems of changing the specification of existing good-types.

New Goods Bias

A failure to pick up new products in a timely manner will reduce the ability of the RPI to cover an average consumption basket, and consequently its ability to measure prices accurately. Basic consumer theory suggests that new goods should be incorporated into an ideal price index by assuming that their price prior to appearance was equal to the reservation price. In this case, as demand rises above zero (i.e. the item features in expenditure surveys), the price will fall below the reservation price. Consequently, a Laspeyres index, such as the RPI, which is based on expenditure weights from the previous period, will overstate inflation. The following example clarifies the issue.²

Table 2 **New goods bias in Laspeyres indices**

Period	p_{1t}	x_{1t}	p_{2t}	x_{2t}	Laspeyres	Paasche	Fisher
0	1	1	4	0	100.0	100.0	100.0
1	1	0.9	2	0.05	100.0	90.9	95.3

where p_{it} refers to the price of good i at time t , x_{it} refers to the quantity purchased, and the budget constraint is one unit value. Good 2 is a new good, introduced in period 1. Thus, its period 0 price is set at the reservation price (4) and its quantity is zero. In period 1, 0.05 units of good 2 are purchased at a price of 2 per unit - 10% of total expenditure. The Laspeyres index (the base form used in construction of consumer price indices) uses weights based on expenditure in the previous period. Consequently, in period 1, the Laspeyres index ignores good 2 and thus fails to capture the 'fall' in price from reservation to actual. By contrast, the Paasche index (using end-period

² This example is a variant on one used by Oulton (1995).

weights) and the Fisher index (which is the geometric mean of Laspeyres and Paasche) reflect the notional fall.

A failure to incorporate new goods may lead to bias of two forms. First, the failure to price any good at its reservation price prior to inclusion will lead to an upward bias. However, this is not a practical concept for the macro-economist. Over a twelve month period, a bewildering number of new brands and products will become available - according to Hausman (1994), around 190 new brands of cereals became available in the US between 1980 and 1992! And if one considers the locational and 'state-of-the-world' characteristics of any good in addition to purely physical characteristics, the set of new goods is potentially infinite. For example, an ice cream on a sunny day is presumably a new good when all previous history had been cloudy. Since reservation prices are calculated relative to other prices, their use is not viable if there are too many new goods for which prices must be derived. So at a practical level, only the second source of bias is a tractable issue for the macro-economist. A failure to introduce newly available goods to the index promptly may cause further upward bias, if the price paid for those goods falls relative to the RPI following introduction.

At a practical level, new goods bias refers to any overstatement of inflation due to the delayed introduction of a good which has been on the market for some time. Note that such a delay need not imply systematic upward bias. For that to occur, a subsidiary hypothesis must hold. A lagged introduction of a new good will only lead to an overstatement of inflation if the price of that good had been falling relative to the price index. If goods follow a pricing cycle on their introduction, with relative prices falling from an initially high level, a failure to incorporate them swiftly will lead to an overstatement of

inflation. The extent of any new goods bias depends on both the decrease in the relative price of new goods and the share of total consumption accounted for by new goods excluded from the price index.

Quality Adjustment Bias

As with new goods bias, this problem is rooted in the difficulties faced when constructing the consumption set underlying the price index. The use of a fixed consumption set assumes implicitly that the nature of goods bought does not change over time. That is, the specification of goods within a given category does not change. In practise, the specification of goods does change over time, so that data-collectors may be unable to find a price-quote for a good of the precise form that they had previously been sampling. In this case, a product with a different specification must be incorporated into the price index. As the specification of products changes, so does the set of product characteristics and features - in other words, quality changes. In the very long run, competitive forces should ensure that any relative price change is due to quality. But equivalence of quality-adjusted prices need not occur in the short run. In this case, any failure to account for the difference between the set of services offered by the old and new specifications (i.e. differences in quality) will lead to inaccuracy in the price index. If quality is seen as increasing over time (perhaps due to technological advance), a failure to deal with it will bias the RPI above a notional 'true cost of living index', overstating inflation.

In the case of consumer durables, the change in quality may reflect either a difference in utility derived per period or of the length of the period over which the flow of utility is obtained. Making clothing more fashionable may improve quality on the first count, but reduce it on the second.

In practise, all statistical agencies make some attempt to control for quality changes. When a data-collector can no longer obtain a price quotation for a given specification (for example because the store no longer stocks that brand), a new good is substituted for the old. Both the Bureau of Labor Statistics (US) and the Central Statistical Office (UK) use a fairly rudimentary technique to adjust for any quality changes - sometimes termed the matched models method. Following a specification change, one of a number of assumptions concerning quality changes can be imposed:

- 1) the prices of the old and new brands may be 'linked', so that any difference in price between the two is assumed to be due to quality differences.
- 2) zero quality change may be assumed, so that the prices of old and new varieties are compared directly.
- 3) it may be assumed that the substitute products' price has risen in line with other products in the same sector. In this case, the item may be omitted for the month of the specification change.
- 4) some attempt may be made to measure quality change (usually in terms of changes to producer production costs).

Bias will only exist to the extent that attempts to adjust for quality change fail.

Substitution Bias

In common with other consumer price indices, the UK's Retail Prices Index measures a weighted average of the prices paid by consumers for a fixed basket of goods and services. Individual constituent parts of the RPI are weighted according to the share of expenditure on the item - as estimated in the previous year's Family Expenditure Survey. Consequently, the RPI is based on weights calculated before price changes are revealed - it is a form of Laspeyres index.

Since it is based on a pre-fixed basket of goods, a Laspeyres index cannot pick up any incentive for consumers to substitute away from relatively more expensive items. Consequently, it can only measure inflation in the cost of living accurately when preferences are Leontief (i.e. agents do not substitute in response to relative price changes). Ignoring the capacity of consumers to alter their spending patterns when relative prices change leads Laspeyres indices to overstate inflation. The extent of any substitution bias depends on both the extent to which households substitute between goods in response to relative price changes, and the extent to which relative prices change during the period in which weights are fixed. Note that this bias is *a priori* upward - it arises from a failure to match the optimising agents' expenditure patterns. Since these will minimise the cost of a given level of utility, any divergence will 'raise' the recorded cost of living.³

³ Note that in the [implausible] instance of an index dominated by Giffen goods, the failure to allow for substitution effects will cause a downward bias.

There are two possible aspects of substitution bias: product (or output) bias and outlet bias. Product bias arises to the extent that the price index fails to pick up changes in the expenditure shares of goods within the basket of goods bought by consumers. Outlet substitution bias refers to any failure to capture changes in the locational characteristics of the basket - a shift towards lower-priced outlets will lead to an overstatement of inflation by a Laspeyres index.

Product Bias

Product bias is rooted in the fixed nature of the weights by item. By failing to capture the evolution of expenditure patterns over time and state, price indices do not cover the cost of living perfectly.

Outlet Substitution Bias

In addition to weights by item type, the price observations are also (implicitly) weighted by store type. The implicit weights are due to the number of price observations taken at different locations. Since a prices index will be based on a fixed locational sample, and optimising agents may shift location in response to relative price movements, there is scope for bias. A fixed-weight index will not pick up changes in shopping habits. Outlet substitution bias refers to the difference between a fixed store-weight index and a cost-of-living index.

Within a sector of the prices index, outlet substitution bias will occur if price changes at uncovered stores differ systematically from those in covered stores (after adjustment for quality differences). Note that a 1-off change in price levels will not affect the index. For example, if uncovered stores always price

10% below established stores then an index calculated without them will be identical to one which includes all stores - regardless of changes in market share. This is a counter-intuitive result, but one which follows from both the similarity of Laspeyres and Paasche indices when relative prices are constant and the nature of the 'ideal' cost-of-living index.

Laspeyres and Paasche indices are identical when relative prices do not change.

This result may be demonstrated by an algebraic comparison of Laspeyres and Paasche indices.

Let there be two goods (a and b) characterised by the store in which they are sold. In this case, the equations below describe the Paasche and Laspeyres index formulation in period $t+1$. The Laspeyres index is based on initial period weights, while the Paasche index is based on current period weights.

$$P_L \equiv \frac{P_2}{P_1} \equiv \frac{p_{a,t+1}x_{a,t} + p_{b,t+1}x_{b,t}}{p_{a,t}x_{a,t} + p_{b,t}x_{b,t}} \quad (1)$$

$$P_P \equiv \frac{P_2}{P_1} \equiv \frac{p_{a,t+1}x_{a,t+1} + p_{b,t+1}x_{b,t+1}}{p_{a,t}x_{a,t+1} + p_{b,t}x_{b,t+1}} \quad (2)$$

Let the prices of the two goods evolve such that relative prices are unchanged
- they are simply scaled by the same factor.

$$p_{a,t+1} = \alpha p_{a,t} , p_{b,t+1} = \alpha p_{b,t} \quad (3)$$

To show that market share changes have no effect, let the goods shares be scaled as follows

$$x_{a,t+1} = (1 - \beta)x_{a,t} , x_{b,t+1} = \beta x_{b,t} , \beta > 0 , \beta < 1 \quad (4)$$

In this case, the Laspeyres index is

$$P_L = \frac{\alpha p_{a,t} x_{a,t} + \alpha p_{b,t} x_{b,t}}{p_{a,t} x_{a,t} + p_{b,t} x_{b,t}} = \alpha \quad (5)$$

and the Paasche index is

$$P_P = \frac{\alpha p_{a,t} (1 - \beta)x_{a,t} + \alpha p_{b,t} \beta x_{b,t}}{p_{a,t} (1 - \beta)x_{a,t} + p_{b,t} \beta x_{b,t}} = \alpha \quad (6)$$

The two are identical. This result is due to the constant relative prices. A price index measures the rate of change of prices. If all prices are changing at the same rate, then the weight attached to each good is irrelevant. But if relative prices do change then both the rate of change of relative prices and the change in market share are important.

Let the price of the two goods evolve as follows

$$p_{a,t+1} = \alpha p_{a,t} , p_{b,t+1} = \delta \alpha p_{b,t} \quad (7)$$

Market share evolves as in (4) above.

In this case, the Laspeyres is

$$P_L = \frac{\alpha p_{a,t} x_{a,t} + \delta \alpha p_{b,t} x_{b,t}}{p_{a,t} x_{a,t} + p_{b,t} x_{b,t}} = \alpha + \frac{(\delta - 1) p_{b,t} x_{b,t}}{p_{a,t} x_{a,t} + p_{b,t} x_{b,t}} \quad (8)$$

and the Paasche is

$$P_P = \frac{\alpha p_{a,t} (1 - \beta) x_{a,t} + \delta \alpha p_{b,t} \beta x_{b,t}}{p_{a,t} (1 - \beta) x_{a,t} + p_{b,t} \beta x_{b,t}} = \alpha + \frac{(\delta - 1) p_{b,t} \beta x_{b,t}}{p_{a,t} (1 - \beta) x_{a,t} + p_{b,t} \beta x_{b,t}} \quad (9)$$

The two differ - the extent of the difference is governed by the scale of the relative price differential and the extent of market share shift.

However, this algebra ignores potential switchover effects due to search behaviour. For example, suppose there are three agents in the economy. In period 1, two agents buy all their goods at shop A and 1 at shop B, which is more expensive, but of similar quality. In period 2 neither shop changes the prices offered, but the third agent switches and also purchases all goods from shop A. Clearly, prices paid have fallen on average (since the third agent pays less and the others pay the same). But neither the Laspeyres index (which is weighted pre-switch) nor the Paasche (post-switch) will record any price change. This is essentially a problem of parallel markets (see Silver (1989) for a more detailed exposition).

Nevertheless, this manifestation of the parallel markets issue need not cause the RPI to deviate from an 'ideal' cost-of-living index. The 'ideal' index was defined to reflect utility-maximising decisions. To be more precise, in a world of uncertainty, the ideal index may be seen as reflecting maximisation of expected utility, or *ex ante* utility maximisation. And any switching behavior exhibited by the third agent may be viewed as an *ex post* response to search behavior. In this case, the only reason for the switch is that an imperfectly

informed agent has, through costly search, found a cheaper store. In this case, realised prices have fallen. But *ex ante* the cost of achieving a constant level of expected utility was unchanged.

In practical terms, the extent of outlet bias will vary according to differentials in quality-adjusted price inflation between store-types; and shifts in market share between store-types, relative to changes in the weights attributed to them within the RPI. This factor depends on the flexibility of 'store-type' weights within the RPI, and consequently the CSO's methodology.

3. Compositional Bias - Quality Bias

Gordon (1990) has provided the most detailed assessment of quality bias in the measurement of durable goods prices. In addition to an analysis of durables in the producer price indices, Gordon used catalogue data to assess quality adjustment bias amongst selected categories of consumer durables. He concluded that the US Consumer Price Index overstated the rate of increase in durables prices by an average of 1.54% pa between 1947 and 1983. However, the errors appear to have declined over time - between 1973 and 1983, the average overstatement was 1.05% pa.

Gordon's estimates have provided a basis for extrapolation in Canada and the United States. Lebow *et al* (1992) extrapolate from the 1.5% pa figure across the 22.56% of the CPI sample accounted for by durables.⁴ This yields an estimate of quality-adjustment bias of around 0.3% pa. Crawford (1993) also extrapolates from Gordon's figures, to estimate a bias of approximately 0.2% pa in Canada (he uses the 1% pa bias in consumer durables). The accuracy of these guesstimates depends on both the accuracy of Gordon's study and the applicability of the extrapolations carried out.

⁴ The sectors covered are: housefurnishings - appliances, housefurnishings - lawn equipment and power tools, apparel commodities, new and used vehicles, airline fares, medical care commodities and services, and entertainment commodities.

Scope for Extrapolation

Most work to date has focused on quality adjustment in durables. However, while quality-change in durables is easier to visualise, it should not be ignored in the non-durable sector. For example, a move towards eco-friendly washing powder is a quality change, and should be treated accordingly. In this case, it is tempting to extend the range of extrapolation. However, there are no *a priori* reasons to expect quality bias in nondurables to be of the same direction, let alone the same extent as the bias found in consumer durables prices.

In his book, Gordon notes that one cannot extrapolate his figures across all official price indices: "In particular, there is no necessary conflict between the new results and the long-standing claim by Triplett ... that the overall bias in official price indexes is just as likely to be downward as upward". His point is that, in the absence of data on consumer non-durables or services, it is not possible to be at all sure about the extent of bias in the CPI. Lebow *et al* and Crawford assume that bias in consumer nondurables and services is as likely to be downward as upward (i.e. nets out at zero).

Extrapolation Across Countries

While in principle, there is no reason to expect the Bureau of Labor Statistics (BLS) methods (usually the matched models method) to lead to systematic bias, Gordon found that they did. Since the treatment of quality is essentially similar in the United States and the United Kingdom, one might use Gordon's research as the basis for a guesstimate of bias in the UK. However, any estimates should be treated with considerable caution, given the uncertain

theoretical base for the bias. Gordon's 'true indices' were constructed using a similar technique to the BLS - the matched models method. Consequently, the US bias appears to be more administrative than methodological. In this case, extrapolation to the UK will only be valid if the UK administration suffers from similar flaws. Oulton (1995) did not find any *a priori* reasons to expect the CSO to perform better than the BLS, which may justify the use of Gordon's estimates as a rough guide to UK bias.

Accuracy of Gordon's Estimates

Data Problems

Gordon's study is based on catalogue data from trade magazines and retail catalogues. Triplett (1993) notes that the data does not correspond exactly with the BLS sources, suggesting a margin for error. For example, when Gordon analyses the 'construction machinery' component of the PPI he is forced to compare catalogue data on post-hole diggers with the PPI data for trenching machines, because the Sears catalogue does not contain the latter. However, while the potential for inconsistencies between the catalogues and the CPI sample leave a margin for error, it is of a random direction. Consequently, there are no *a priori* reasons to prefer a lower or higher figure than Gordon's estimate.

Problems With the Matched Models Method

In line with current BLS practise, Gordon uses the matched models method to estimate his price indices. However, it is unclear how sensitive this technique is in areas of rapid technological change. Where change is continuous, a

binary system - in which price changes are either recorded as quality changes or not - may not give an accurate assessment of the extent of any quality change. Given the pace of technological change in some areas (such as home computers), Gordon's matched models estimate may be too low.

An alternative estimation procedure focuses more closely on the quality concept. Hedonic regressions are sometimes used for goods which are subject to frequent quality adjustments. Here, the price of the good is viewed as the sum of prices paid for its various characteristics. For example the price of a camera may be viewed as the sum of the price paid for its basic functions, lens capabilities, bulk, brand reputation, and a residual of other factors:

$$PRICE = \alpha[BASIC]_t + \beta[LENS]_t + \gamma[BULK]_t + \delta[BRAND]_t + \varepsilon_t$$

Berndt and Griliches (1993) estimated hedonic regressions for personal computers available between 1982 and 1988. Their work (which extends earlier studies by two of their students) examines a number of product characteristics including the amount of hard disk storage; the clock speed and the amount of random access memory on each machine. Table 3 (overleaf) compares Berndt and Griliche's price index with earlier hedonic indices and both Gordon's estimate and the official computer price index. All are expressed relative to the all-items CPI. The two key points to arise from the table are that the official computer price index did capture much of the quality-adjusted fall in prices and that Gordon's matched models technique understates bias in this sector by approximately 2pp pa.

Although there are a number of caveats, current best-practise hedonic regressions suggest that Gordon underestimates bias in this sector by 2pp pa. However, even if Gordon understates bias in a few goods, the overall impact

need not be very significant. The Berndt *et al* results are applicable to just one area of rapid technological advance. If this extra bias were extended to all electrical appliances (as a proxy for high tech goods), the extra bias in the index would be minimal - electrical appliances make up just 1% of the RPI so that an extra 2pp bias pa would only lead to 0.02pp pa bias in the overall index.

Table 3 **Alternative measures of quality-adjusted computer prices, relative to the CPI***

	Official	Gordon (1990)	Berndt / Griliches (1991)**	Cohen ^a	Kim ^b
1982	1.000	1.000	1.000	1.000	1.000
1983	0.777	0.746	0.448	-.712	0.687
1984	0.568	0.558	0.746	0.593	0.617
1985	0.511	0.387	0.458	0.390	0.409
1986	0.369	0.292	0.282	0.258	0.268
1987	0.321	0.220	0.187	0.191	0.194
Average annual growth rate (1982 - 1987)	-20.33%	-26.12%	-28.49%	-28.16%	-27.96%

** Time and Age specification, allowing for discounts from list price

a Doctoral thesis

b Doctoral thesis

*Source: Berndt and Griliches (1993)

However, the technological effect need not be constrained to the new 'high tech' products, so that if there are significant disparities, they may well be widespread. In a recent paper, Nordhaus (1994) examined the measurement of the price of light over time, to find a significant disparity between a 'true' index and actual price measures (the former rose between four and seventy-five times as far as the latter). If his higher estimate is correct, then the quality bias for light was some 3.6% pa, more than double Gordon's figure for

bias in consumer durables prices (but less than the bias found in the computer sector).

A key difference between Nordhaus' and Gordon's work is Nordhaus' focus on the flow of services aspect of durable goods consumption. Rather than analysing the price of light bulbs and electricity, he assesses the price of light. Although an assessment of bias between the RPI and a cost of living index incorporating the flow of services element of durables consumption would be ideal for our purposes, it is not currently tractable for the index as a whole. This is because, for many commodities, it is hard to define, let alone measure, the flow of services being consumed. For example, it is hard to see what services we might price when evaluating the cost of music. Indeed, even light is not clearly a *final consumption good*. We might prefer a measure of the cost of being able to carry out activities for which light is an input (such as reading at night). Since we are unable to construct estimates across the entire RPI, we offer a range for quality bias excluding any service-flow aspects.

Our results should be treated as an estimate of just part of the total possible bias. The limited evidence of Nordhaus' study suggests that total bias might be substantially greater (extrapolating across all durables, using his 3.6% figure suggests quality bias of up to 0.8pp pa). However, extrapolating from just one study is risky. Furthermore, there is some reason to expect bias in the *recent* measurement of the price of light to have been less than that estimated by Nordhaus.

Nordhaus uses a fairly broad approach, to estimate average bias over a long run. His study suggests that estimates of real wage growth over a century or more are significantly understated, because they do not allow for

technological change. The 3.6% figure may be considered as a mean bias over the full 1800 to 1992 sample. If quality bias has remained broadly constant over time, then the 3.6% figure might cast some doubt on Gordon's results. However, there is some evidence that quality bias has declined over time - perhaps reflecting improvements in data collection methodology. Gordon found quality bias in consumer durables to have averaged 2.21% pa in the 1947 and 1960 sub-period, but just 1.05% in the 1973 to 1983 period. If quality bias has fallen over time, then a 3.6% two century mean may not be too different from Gordon's estimate for recent years.

Not all criticisms of Gordon's work suggest that he under-estimates bias. In his review of Gordon's book, Triplett (1993) notes that Gordon rejects his hedonic car price index in favour of a matched models approach. While there are some reservations about the hedonic index, "which is probably biased in the opposite direction from the CPI",⁵ its exclusion is not consistent with general practise throughout Gordon's work. Substituting the hedonic index for the reported index reduces the overall durables bias estimate by around a third.

There is also some evidence to suggest that quality bias has been negative in some sectors. Liegey's (1993) study of apparel indexes in the US found that the CPI tended to understate price increases. And research by Statistics Sweden also suggests a negative bias in the clothing and footwear sectors. The downward bias may be seen as a product of the seasonal nature of the clothing series. At the start of each season, fashion items are introduced. Towards the

⁵ Triplett (1993), page 244.

end of the season, their price tends to be reduced, prior to the release of the next season's fashions at higher prices. The bias arises because, as the season progresses, the goods become less fashionable. This is an intangible quality fall that is not excluded from the price measurement because the old brand remains available. When the new clothes are released, the change in quality is accommodated, because a new brand has replaced the old. So there is no offsetting upward bias at the start of each new season. Consequently, we might expect a downward bias in sectors which are subject to seasonal fluctuation in fashion.

Gordon's estimate provides a useful basis for extrapolation - it is possibly the best available estimate of bias between the US CPI and an acquisition cost of living index. But the additional bias arising from the flow of services should be considered when assessing any estimate of quality bias.

Application to the UK

In the absence of empirical research specific to the UK, our strategy is to impose assumptions based on work in the US on one fifth of the RPI sample spent on durables.⁶ Imposing Gordon's estimate of US bias yields a range of 0.2 to 0.3% pa upward bias to the RPI. Since there are no *a priori* reasons for preferring either end of the range, we use its mid-point as a most likely bias: roughly 0.25% pa. However, this estimate is sensitive to both the proportion of the RPI basket deemed eligible and the bias estimates imposed.

⁶ The Household goods, clothing and footwear, and leisure goods sectors.

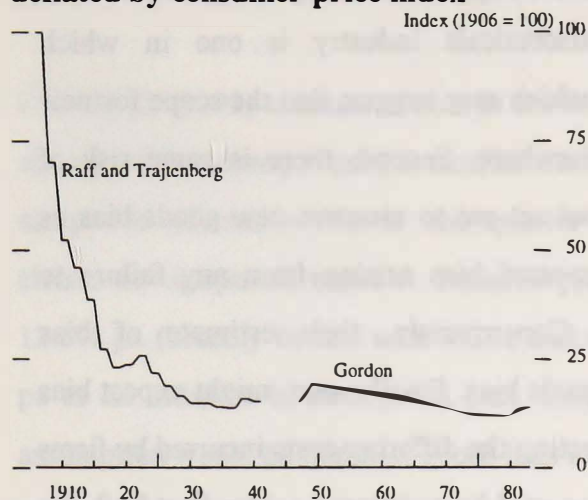
The *circa* 6% of the RPI sample accounted for by 'purchase of motor vehicles' is not included in the sample of durables liable to quality adjustment. If cars are included, the range runs from 0.25 to 0.35% pa. The section is omitted because the peculiar nature of the component's construction does not leave it open to systematic bias in the same way as other components are. The motor vehicle purchase component of the RPI is derived entirely from the price changes of two and three-year old used cars, as quoted in Glass's Guide. New car prices are not included, because of problems with quality adjustment, obtaining transactions rather than list prices, and the desire to exclude business purchases from the RPI. Although there is some debate about the validity of used car prices as a proxy for new car prices, a beneficial spin-off of this technique is that the car price component is less susceptible to quality bias than other sections. Since individual models of the age in question do not tend to drop out of the used car market during the year, the used car sample need only be adjusted once a year. Consequently, the problem of linking new to discontinued models does not arise to the same extent as it does for other items. Furthermore, there is more reason to expect any price differences to reflect quality. It is hard to argue that competitive forces will not have led to quality-adjusted price equivalence after two to three years.

4. Compositional Bias - New Goods Bias

New goods bias refers to discrepancies due to the failure to incorporate entirely new product types. This is distinct from quality bias, which is due to problems incorporating new brands of existing products. In theory, the failure to price new goods at their reservation price prior to inclusion may lead to biases. But, given the plethora of new goods becoming available, assessment of this theoretical bias may not be practicable (see Section 2). However, new goods can lead to bias if they are available (i.e. not priced at the reservation price) and yet not incorporated into the price index. We focus on this area. The level of bias is dependent upon the proportion of consumer expenditure on new products and the extent of any pricing cycle they follow. Raff and Trajtenberg (1995) estimated a quality-adjusted price index for automobiles between 1906 and 1940 - years when the automobile was a 'new good'. As Chart 1 (overleaf) shows, the bulk of the price fall occurred during the first few years of the product's introduction - there is clear evidence of a pricing cycle. It is interesting to note that since 1920 there has been little decline in quality-adjusted price, despite multiple specification changes.

Chart 1

Quality-adjusted automobile price index deflated by consumer price index*



* Source : Raff and Trajtenberg (1995)

Berndt, Griliches and Rosett (1993) found a bias of some 3% pa in the prescription pharmaceutical preparations component of the US Producer Price Index. While the official price index for this component rose at an average annual rate of 9.09% from January 1984 to December 1989, a Divisia index based on sales data for all prescription pharmaceutical products sold by 4 firms rose at just 6.03%. Assuming that the firms were representative, this implies a bias of 3% pa. The authors found the 4 firms for which they had data to be representative of the BLS coverage (the 4 firms accounted for 24% of total industry sales, and the average annual price increase of items sampled by the BLS from the four firms was 8.94%). Further analysis showed that the BLS oversample middle-aged products at the expense of new products (under two years old) and that the price of new products rose more slowly than average. Consequently, the bias can be attributed to a failure to accommodate new goods adequately.

Although Berndt *et al* analysed a component of the PPI, biases might be expected in a measure of consumer prices. In their sample, the PPI errors derive from under-sampling products aged two years or less. It seems unlikely that the CPI coverage of such products would be more representative. However, using the Berndt *et al* figures as a basis for estimation is risky for several reasons. First, the pharmaceuticals industry is one in which technological change is significant, which may suggest that the scope for new goods bias is greater there than elsewhere. Second, there is some risk of double-counting. Berndt *et al* did not set out to measure new goods bias as distinct from other sources, and assessed bias arising from any failure to adjust for specification changes. Consequently, their estimate of bias incorporates both quality and new goods bias. Finally, one might expect bias to vary along the supply chain - reflecting the differing costs incurred by firms at each stage of production. The price paid by consumers will reflect both the price paid by wholesalers and other input costs (such as labour). Since these further unit costs might be expected to be identical for both old and new goods, there is no reason to expect them to contribute towards bias. In this case, one might expect less bias in the consumer price index, since a smaller proportion of the price paid by consumers is open to bias.

Studies of bias in the American and Canadian Consumer Price Indices have taken a more general approach. Bias is taken to be equal to the product of the rate of decline of prices of new goods (relative to the index) and the proportion of the index open to new goods bias.

In their estimate of bias in the United States, Lebow *et al* (1992) assume that new goods bias is restricted to 2.4% of the CPI sample.⁷ The average shift in relative prices is assumed to be 20% pa - "about the same rate of relative decline posted by prices for computing equipment"(page 20). From this, they derive an average annual bias of approximately 0.5%.

Crawford's (1993) estimate of bias in the Canadian CPI is considerably lower, although the technique used is similar. He assumes that just 0.5% of the CPI sample is affected - that is one quarter of the household appliance and electronic equipment sections. Relative prices are assumed to fall by some 12.5% pa (broadly in line with VCRs and microwave ovens), to yield 0.06% pa as an estimate of new goods bias. Consequently, "a reasonable working assumption would be that new goods bias in the Canadian CPI is less than 0.1 per cent per year" (page 54).

One potential problem with both the studies above is that they are restricted to a small part of the CPI basket. The new goods bias problem is assumed to arise only in more technologically advanced sectors. However, new entrants may come from a wide range of sectors - in 1994 'leggings' and fresh peaches were amongst the new entrants to the United Kingdom's RPI, in addition to the 'technological' introduction of camcorders. However, while the estimates ignore many new goods, this may be reasonable if relative price movements are not significant.

⁷ The sectors assessed are: housefurnishings - appliances, housefurnishings - lawn equipment and power tools, and medical care commodities.

About 30 items were added to the UK's RPI in 1994. But excluding these products does not have a systematic effect on the section-level price-indices concerned. The section most affected was 'soft drinks' - excluding new entrants reduces the index by 0.7 points in March 1994, but raises it by 1.6 points in October.⁸ Since the effect is not systematic, ignoring new goods where there is no *a priori* expectation of relative price movements may prove valid. In this case, the reduced samples of Crawford and Lebow *et al* may be justified. A key distinction appears to be between new entrants which have only been available for a short time (such as camcorders) and new entrants which have been available for a considerable period, (such as fresh peaches). There may be *a priori* reasons to expect a pricing cycle (and hence bias) in the former, but not in the latter.

The considerable difference between the US and Canadian estimates does not represent a fundamental difference in index construction methodology so much as differing judgements by the authors concerning both relative price movements and the extent of new goods in the CPI. Since the estimates were based on a similar technique and there are no *a priori* reasons for new goods bias to be five times greater in the United States than in Canada, one set of judgements must be better than the other.

Crawford's assessment of the proportion of the basket exposed to bias is more plausible, because he allows only part of each section to be affected in each period. Lebow *et al* assume that all goods in any section of the CPI open to bias will have their prices overstated. But this is not possible - a sectors'

⁸ Source: CSO unpublished research.

existence ensures some recording of prices, so that only a part of the purchases due to each sector can be unrecorded. Crawford assumes that bias is applicable to $\frac{1}{4}$ of each sector considered. This is probably an upper estimate in categories such as 'electrical appliances' which apply to many goods, only a few of which are unrecorded in any year.

The relative price assumptions used are relatively arbitrary. Since neither is clearly preferable, it may be better to analyse a range of new goods bias possibilities - using Crawford's proportion of the index sample, and both authors' assumptions concerning relative price movements. For an upper limit, imposing Lebow *et al's* 20% pa on Crawford's 0.06% of the RPI sample yields 0.1% pa as an estimate of new goods bias. The lower limit is 0.06% as in Crawford's paper.

Application to the UK

The RPI sample is revised annually. In addition to weighting changes, items are admitted to and deleted from the sample - about 50 changes are made each year, although in many cases, the revisions are merely a change in specification - for example from normal to low fat yoghurt. Although the sample revisions are based on an analysis of the Family Expenditure Survey, there are no scientific criteria for selection.

There may be significant lags prior to the inclusion of a new product. For example, video recorders, CD players and microwaves were not included until 1987 - by which time some 43.5% of households owned a video-recorder. In Canada, video recorders were added to the CPI in 1985. However, CD players were not added to the Canadian CPI until 1990 - later than in the UK.

Consequently, there are no firm *a priori* reasons to expect the lag between the widespread consumption of a good and its inclusion in the price index sample to be smaller (or larger) in the United Kingdom than elsewhere. In this case, one might expect biases of a similar magnitude to those estimated in the United States and Canada - imposing Lebow's 20% fall in relative prices on one quarter of the electrical appliance and audio-visual components of the RPI yields an estimate of bias of 0.126% pa ($0.2 \times 0.25 \times 2.1 = 0.126\%$ pa).

To test the robustness of this simplistic estimate, several simulations were run. A set of new goods is assumed to enter the market place in January 1987, but not to enter the RPI sample until 1991 (this date can be varied by several years without altering the overall results significantly). By varying assumptions about the relative price movements of the new goods and the 'share' of each sector they account for, we derive a range of plausible estimates of new goods bias. The simulation is restricted to electrical appliances and audio-visual products - to reflect the narrow range of goods for which an *a priori* pricing cycle might be expected.

We make 3 assumptions about price movements of the new goods relative to the RPI - a 12% fall per annum, in line with Crawford's estimate; a 20% fall per annum, in line with Lebow's estimate and the CPI computing component; and a 30% pa fall, in line with Berndt *et al's* quality-adjusted computer-price estimates (recorded in Table 3). Although computers have followed a notable pricing cycle, this is a reasonable imposition, given the restriction of the 'bias sample' to electrical appliances and audio visual goods. However, it might be treated as an upper bound, since relative price declines appear to have been greater for computers than for other new goods. We also make 4 assumptions about the eventual share of the sector taken up - 25%, 20%, 10%, and 50%.

The 50% assumption is included solely as an extreme reference point. The estimates of bias did not prove very sensitive to shifting from 10% to 25% sector shares.

The results of the simulations are shown in Table 4. A 'plausible range' for new goods bias runs from 0.02 to 0.16pp per annum.

Table 4 **New goods scenario results**

rel P changes	New goods bias - % per annum			
	share = 10%	share = 20%	share = 25%	share = 50%
-12% pa	0.02	0.03	0.04	0.06
-20% pa	0.07	0.09	0.10	0.15
-30% pa	0.12	0.15	0.16	0.24

5. Substitution Bias - Product Bias

Most early work on systematic bias focused on this source. The tests have followed one of two forms - comparing the actual price index with a 'superlative' index (more akin to an ideal cost of living index), and comparing a Laspeyres price index to an estimated cost of constant utility index.

Under the superlative index technique, the actual CPI is compared with a superlative index. These indices use weights at the beginning and end of the sample, to derive a price index which corresponds to an ideal cost of living index under a more general set of conditions than either Paasche or Laspeyres indices.⁹ Substitution bias is estimated as the difference between the superlative index and a Laspeyres index of the form used to calculate the CPI.

The alternative approach is 'utility-based', and involves comparison of a Laspeyres price index with a 'true' Cost-of-Living index derived from a set of observed prices and a set of demand functions. These demand functions are derived from actual price observations and an imposed functional form for utility. As a consequence of this imposition, all computed cost-of-living indices are *estimates* of the true index. However, Braithwait (1980) found that his estimates of bias were not sensitive to the functional form imposed, although his conclusion is "tempered by the restrictiveness of the three models employed" (page 74).

⁹Two popular superlative indices are the Fisher index and the Tornqvist index. The former is the geometric mean of a Paasche and a Laspeyres index - its use derives from the observation that while Laspeyres indices tend to overstate inflation, Paasche indices tend to understate it. The latter is a product of prices ratios, weighted according to expenditure shares in both periods.

There have been a number of studies of product substitution bias in the US. Early studies (reviewed in Triplett (1975)) found that substitution bias was no greater than 0.1% pa, but that it tended to be greatest in periods of high inflation. However, in a later study, Manser and MacDonald (1988) used a more disaggregated data-set, and estimated significantly greater bias. Table 5 summarises the key points of their paper and an earlier paper representative of the results surveyed by Triplett.

Table 5 Substitution bias in the US CPI

	Technique	bias (% pa)	other notes
Manser and MacDonald (1988)	utility-based	0.14 to 0.22	They found that the results were sensitive to the level of aggregation of the data used.
	superlative index	0.18	
Braithwait (1980)	utility-based	0.10	Less detailed study than Manser and MacDonald Found that lengthening the time over which weights were fixed increased bias.

Application to the UK

Since product substitution bias is rooted in the fixity of the index's weights, one would expect to find most bias in indices with the most rigid weighting structure. The more frequent the weighting revisions, the less the scope for the index to miss shifts in consumption patterns, and thus the less the scope for systematic substitution bias. Work by Génèreux (1983) on the Canadian

CPI confirms the sensitivity of bias to the frequency of revisions. He found that substitution bias over the 1957-1978 period could have been reduced from around 0.2% pa to 0.09% pa if the CPI had been re-weighted eight times rather than once (as actually occurred).

Table 6 (below) shows the frequency of weighting revisions within the G7. The UK's RPI is re-weighted annually, while the US CPI is still based on weights calculated between 1982 and 1984. Consequently, one would expect less bias in the UK than in the US¹⁰. Note that, despite the frequent revisions, there is still some scope for bias in the UK's RPI - not only is there rigidity over the course of the year, but the UK's weights are based on data taken from the Family Expenditure Survey of the year prior to re-weighting, so that the January 1995 index was based on expenditure patterns in 1993.

Table 6 Consumer price indices in the G7

Country	Index title	Frequency of weighting - revisions
UK	<i>Retail Prices Index</i>	ANNUAL
Canada	Consumer Prices Index	4 YEARLY
France	Consumer Prices Index	ANNUAL
Germany	Consumer Prices Index	5 YEARLY
Italy	Consumer Prices Index	3 to 4 YEARLY*
Japan	Consumer Prices Index	5 YEARLY
USA	Consumer Prices Index	10 YEARLY

*The Italian weights are not revised according to a rule, but were revised in 1996, 1992, 1989 and 1985.

In many cases, the re-weighting frequencies do not follow a rigid rule. For example, German re-weighting broke trend to avoid coinciding with unification. The Canadian revision was put back from 1993 to 1995, to allow the weights to adjust during a period of calm, after households had established firm spending patterns following the introduction of GST in 1991.

¹⁰ It is perhaps surprising that, given its more formal objective to measure the cost-of-living, that the US index is likely to be a worse approximation than that of the UK.

In the light of the relatively frequent revisions, a reasonable guesstimate for the UK would be to impose Manser and MacDonald's upper limit for bias in an annually revised Laspeyres index - i.e. 0.05pp per annum. However, this may prove an underestimate - while their study is based on a more disaggregated sample than earlier work, some aggregation was involved in their 101 commodities categories. Since they found higher-level aggregation to understate any bias, it is not unreasonable to assume that the aggregation in their study leads to under-estimation - in other words 0.05pp can be treated as a lower bound. Since there is likely to be some premium on realignment the upper bound will probably be less than that found for the US. G  n  reux's (1983) work on the Canadian CPI suggests that bias is approximately halved when the frequency of revision is raised eightfold. Assuming that this applies to the UK and noting that the UK's RPI is revised ten times more frequently than the US CPI gives an upper bound of around 0.1pp per annum.

6. Substitution Bias - Outlet Bias

There has been less empirical work devoted to estimation of outlet substitution bias - largely because of the increased data requirements. The most detailed work to date is due to Reinsdorf (1993), who used two methods to assess the extent of outlet substitution bias in the United States' Consumer Price Index.

His first technique involves a comparison of price levels in outgoing and incoming CPI outlet samples. Since the probability of an outlet being selected as part of the US sample is proportional to its market share at the time the sample is set up, a comparison of new and old samples will reflect the evolution of consumer outlet choices over the preceding five years. Consumer switches and entry of cut-price stores should both be reflected in a lower mean price amongst the new stores than amongst the old. The difference in prices may be used as an (upper) estimate of bias due to the delayed incorporation of new outlets.

The alternative technique used by Reinsdorf involves a comparison of the evolution over time of unlinked sample average prices and their linked CPI component index counterparts. Both series are calculated using data from a similar set of outlets. However, while the average price series incorporates new outlets without linking (i.e. assumes that price differences are not due to quality), the calculation of the CPI series involves a linking procedure which implicitly assumes zero price dispersion (a chain-linking procedure is used, so that at the time of linking, any price differentials are taken to be due to quality). Since one measure implicitly assumes zero quality change and the other zero price dispersion, a comparison of the 2 will yield an upper bound for price dispersion. Table 7 summarises the results of the two techniques :

Table 7 Reinsdorf's estimates of the effect of outlet price differentials on the US CPI.

		Implied outlet bias (pp pa)
Technique i	Food at home	0.25
	Petrol	0.25
Technique ii	Food at home	2.0
	Petrol	0.9

Both techniques are likely to overestimate outlet substitution bias a little. The first technique includes some effects of product substitution bias. A comparison of old and new CPI outlet samples is not a pure test of the effects of consumer outlet substitution. New samples of items and brands are necessarily drawn at the same time that new outlet samples are drawn. Consequently, for many goods, brand and variety substitution effects will be combined with pure outlet substitution. Since most estimates of total US CPI bias involve summing an extrapolation from Reinsdorf and an estimate of product substitution bias, there is scope for double counting.

The second technique yields an upper bound, and is also open to problems of double counting. By assuming that any difference between the CPI and the average price index is a bias in the CPI (i.e. that there is no difference in quality between old and new outlets) this technique yields an upper bound for bias. Again, the potential for changes in item and brand choice along with location suggests that some double counting will occur if this technique is used to estimate outlet substitution bias alone.

Furthermore, Reinsdorf's figures may also include a 'functional form bias' peculiar to the techniques used by the BLS to rotate its outlet sample. When

the outlet sample is updated by the BLS, the old and new samples are linked together to form a chained Laspeyres index for each good. This chaining can create an upward bias in the price indices for individual components when relative prices cycle up and down. See Moulton (1993) for a more detailed discussion. Coupled with the earlier criticism, this suggests that extrapolation from Reinsdorf is likely to overstate outlet substitution bias.¹¹

On the assumption that some 40% of the CPI basket is open to outlet substitution bias,¹² Lebow *et al* extrapolated from Reinsdorf's estimates to yield a range for total outlet bias of 0.1 to 0.8% pa. However, there are several reasons for viewing the upper part of this range as excessive. Given the potential for double counting, extrapolation from Reinsdorf's higher figures is likely to significantly overestimate outlet substitution bias. Furthermore, to arrive at 0.8% pa bias driven by 40% of the index requires extreme shifts of market share away from currently sampled retailers and extreme shifts in relative (quality-adjusted) prices. The implausible result is probably the result of extrapolating from an upper bound of bias in a 'high-potential bias' sector to a bound for bias across a whole economy.

Crawford's estimate of bias in the Canadian CPI does not extrapolate from Reinsdorf's study. Instead, he imposes plausible figures for relative price and purchase-location shifts to arrive at a rough approximation. Crawford's

¹¹ Note that Reinsdorf did not set out to differentiate between product and outlet bias. Consequently, the criticism above applies more to attempts to extrapolate from his figures than to his own calculations.

¹² The sectors assumed open to bias were: Food and beverages, housing maintenance, household fuels, housefurnishings, apparel commodities, motor fuel, medical care commodities, entertainment commodities, tobacco, and personal care.

guesstimate of outlet substitution bias is the product of the proportion of the RPI basket subject to bias; the average annual shift in market share (for those items subject to the bias); and the average annual difference in quality adjusted relative prices between high and low priced outlets.

Outlet bias is only to be expected for items experiencing shifts in market share between covered and uncovered stores. This clearly excludes any goods provided by a single supplier, such as electricity and water. When forming a rough estimate, it may also be valid to exclude sections for which there is no *a priori* reason to expect a systematic bias in one direction. Crawford assumed that 40% of the basket might be subject to outlet bias.¹³ The sectors chosen are those which have seen the development of warehouse style discount stores. Crawford assumed a 2pp pa shift in market share towards lower priced retailers. This figure is based on a discussion of shifts in market share from independent to chain stores at an aggregate level, and evidence from two large warehouse chains at a microeconomic level. Finally a 10% pa shift in relative prices is imposed as a working hypothesis. Based on these judgements, Crawford arrives at an estimate of outlet substitution bias of 0.08pp pa.

Crawford's figure is significantly lower than the mid-point of the Lebow range. However there are no *a priori* reasons to expect outlet bias to be less in Canada than the United States. While discount stores have a greater market share in the US, it is the growth of market share (rather than its size) which drives outlet bias. The difference between the estimates is rooted in the

¹³ Crawford's sample is fairly similar to Lebow *et al's*, covering food, tobacco, personal care supplies, clothing, consumer durables excluding automobiles, and gasoline.

underlying assumptions rather than any real difference in bias. Lebow *et al* estimate a wide range, the upper end of which seems implausibly large.

Application to the UK

There has been no mechanism for reviewing the sample of stores used in the construction of the United Kingdom's Retail Prices Index - it only changes when an outlet that was being covered either goes out of business or stops stocking the commodities required. When the sample changes, the collectors select 'typical' outlets. There have been no 'scientific' criteria for store-selection, and there has been no mechanism to ensure that existing stores are representative. This failing leaves the collection system open to biases. For example if the market share of a sampled outlet falls due to the appearance of competitors whose price falls relative to the existing stores, the RPI will fail to pick up the fall in price available to consumers, because it covers only the existing store.

The CSO are unsure how rapid the turnover of covered stores is, but it is 'slow'. Our best guess is that the shift in market share toward multiples is under-reported by up to 1 percentage point pa (see Appendix 2 for a review of the method underpinning this guess). Since a failure to review store coverage is essentially the same as a fixed weight system, there is potential for bias in the RPI.

One would expect this bias to be greater than that in the United States, where there is a 'scientific' system of store rotation. The BLS continuously refreshes the outlets within the CPI sample. Approximately one fifth of US cities undergo sample rotation in any year. When a sample is rotated, the new

sample of outlets which submit data is selected by a probabilistic mechanism. This mechanism is applied for each class of goods in the CPI. Each outlet has a probability of selection usually proportional to its share of consumer's expenditure for the good in question, as revealed by data from the Continuing Point of Purchase Survey. Once an outlet has been selected to submit data for a good, the probability of any brand of that good sold in the store being selected is proportional to its sales. At each stage there is a probabilistic criteria for selection of an item for membership of the sample, reducing the scope for systematic error.

Although the American system is 'scientific', there is still scope for outlet substitution bias. Since the outlets are fixed for five years, the CPI essentially involves fixed weights by outlet. In the same way that fixed weights by commodity can lead to bias (product substitution bias) so can fixed weights by outlet. Nevertheless, outlet bias is likely to have been lower in the United States than in the United Kingdom.¹⁴

The RPI outlet sampling procedure has recently been reviewed, as part of the RPI's 'market-testing' process. In the light of this review, some changes are being implemented over the next few years. The introduction of a rolling sample review in conjunction with a probabilistic weighting system should maintain a more representative sample, and consequently reduce the incidence of outlet substitution bias in the RPI.

¹⁴ Note however the scope for functional form bias in the US. This is not inherent in probabilistic rotation, but is a feature of the current BLS approach. Moulton (1993) reviews the issue in depth.

Because the UK's RPI has not been based on a systematically rotated sample and the CSO has not maintained a detailed databank on prices at incoming and outgoing stores, it is not possible to replicate Reinsdorf's work for the United Kingdom. The alternative is to follow a strategy similar to that used by Crawford. We use a series of simulations using plausible inputs, to estimate a likely range for outlet bias and to assess the sensitivity of the estimates. Details are presented in Appendices 1 and 2.

There are 2 possible manifestations of an increase in the market share of lower priced retailers - the expansion of large multiples at the expense of the market share of small independent stores and the expansion of discount retailing in some sectors. Evidence on relative price and market share shifts between independent stores and multiples and discount retailers and established multiples is presented in Appendix 2, and used as the basis for the inputs into the outlet bias simulations. Based on the judgements in Appendix 2, we have run a series of simulations using back data from 1987 to 1994 at the RPI's sector level. The model allows both relative price and market share assumptions to be made.¹⁵

We use 3 relative price scenarios - relative to multiples prices:

- a) Independents rise by 1.875% pa while discount stores fall by 1.875% pa, relative to multiples.

¹⁵ Note that this simple 3-store example is essentially equivalent to an n-store model, where the 3 stores represent the average relative price differentials achieved by the discount and multiple stores. So, for example, a model in which prices charged by independents are diverse but rise by an average of 3.125% relative to multiples will yield similar results to one in which all independents follow the same pricing policy.

- b) Independents rise by 2.5% pa while discount stores fall by 2.5% pa relative to multiples.
- c) Independents rise by 3.125% pa while discount stores fall by 3.125% pa relative to multiples.

Based on the analysis outlined in Appendix 2, scenario b) is probably the most likely. Note that persistent divergence of relative price movements is not a tenable long run solution. But this paper does not seek to estimate bias in the long run. The estimates presented apply to conditions pertinent to the 1990s, and are essentially short to medium term in nature (see Section 7 for a discussion of the possibility of variance in bias across time and state). The relative price scenarios do describe trends apparent in the late 1980s and early 1990s, as supermarkets captured market share from independents and/or defended it against new entrants. Such movements may be expected to be maintained in the near future, as the major players seek to defend their market share. *If these relative price movements do not persist, then this source of bias will clearly become less important.*

We use 4 market share assumptions:

- a) Multiples market share rises by 1pp pa at the expense of independents.
- b) Discount stores market share rises by 1pp pa at the expense of independents.
- c) Multiples market share rises by $\frac{1}{2}$ pp pa, as do discount stores; independents decline by 1pp pa.
- d) Multiples market share rises by 1pp pa, as does discount stores; independents decline by 2pp pa.

Again, based on the analysis outlined in Appendix 2, scenario c) is probably the most likely. A 1pp pa fall in Independents market share is broadly in line with past movements in food retailing (though a little above that elsewhere). But a ½pp pa rise in discount stores share is less than that projected by many analysts (for example James Capel (1994)). Scenario d) is highly implausible, because it implies a fall in independent's market share of approximately double that seen during the 1980s. It is included merely to show that the small range of possible bias (relative to Lebow *et al*) is fairly robust.

The scenarios are evaluated over approximately 35% of the RPI sample - the Food, Tobacco, Clothing, Household goods, and Personal goods and services sub-components. While this selection does not include either DIY or petrol (which might be expected to suffer from bias), this is probably countered by assuming similar relative price movements across all sectors - in practise one would expect smaller movements in the tobacco sector, where price movements are largely excise-driven. Furthermore, some of the sub-sections covered (such as personal services) are unlikely to suffer from outlet bias.

Table 8 summarises the results of the scenarios

Table 8 Scenario results

rel P change	mkt share a)	mkt share b)	Outlet bias - % per annum	
			mkt share c)	mkt share d)
+/- 1.875% pa	0.08	0.15	0.12	0.23
+/- 2.5% pa	0.10	0.20	0.15	0.30
+/- 3.125% pa	0.13	0.25	0.19	0.55

A figure of 0.2% pa is derived using the most plausible set of market share and relative price assumptions. But the data presented is limited to a small

portion of the RPI sample, so it may be preferable to use a range. The bias scenarios range from 0.08 to 0.55% pa. However, the upper estimate is not plausible, because it assumes an excessive fall in Independent's market share. A more reasonable range is 0.08 to 0.25% pa.

The scenario approach appears fairly robust to varied initial assumptions - a comforting result given the scarcity of the data underpinning each one. By exposing the need for extreme market share movements in order to leave this range, the scenarios suggest that outlet bias is unlikely to fall outside it.

7. Variance of Bias Across Time and State

So far, we have couched the discussion in terms of mean bias, using inputs plausible for the 1980s and 1990s. However, there is no reason to presume that any discrepancy between the Retail Prices Index and a hypothetical cost of living index will be constant. There are *a priori* reasons to expect mean bias to have changed over time, and there are *a priori* reasons to expect it to vary according to the state of the economy. An estimation of the fine points of bias variance is beyond the scope of this paper - in this section we merely offer a qualitative assessment of possible shifts in bias which might be anticipated.

Variance over time

One might expect mean systematic bias to have fallen over time, as the technology of index construction improved. For example, the construction of the UK's RPI was improved in 1962, by the introduction of an annual review of weights. And the recent 'market testing' review has led to the implementation of a number of reforms to the outlet sampling procedure, which might be expected to reduce outlet bias. However, the increased proliferation of high tech goods with shortening product cycles may act as an offsetting factor.

Table 9 shows Gordon's estimates of quality bias both across the entire 1947 to 1983 period and during sub-periods. For each component and for the total, quality bias is seen to have fallen over time - possibly reflecting improvements in the practise of index construction.

**Table 9 Quality bias in the US CPI over time:
Gordon's estimate**

	Annual growth rates			
	to 1983 - full sample	1947 to 1960	1960 to 1973	1973 to 1983
Total consumer durables	1.54	2.21	1.24	1.05
Selected components:				
Motor vehicles and parts	1.71	2.39	1.69	0.85
Household equipment and furnishings	1.79	2.52	1.26	1.55

Variance according to state

There are several reasons to expect bias to follow a predictable path around its mean, according to the state of the economy.

Product Substitution Bias

Product substitution bias is driven by the extent of substitution in response to relative price changes and the extent of relative price changes. On average, substitution bias will rise with the time elapsed since the last weighting review - so bias will be greater as the current set of weights come to the end of their life-span.

The extent of substitution in response to relative prices may be expected to be inversely related to income - as agents' incentive to search is higher. There is also some evidence that the magnitude and frequency of relative price changes increases with the level of inflation. A more detailed review of the potential

link between average price inflation and the extent of relative price changes can be found in Fischer (1981). Domberger (1987) found some evidence that relative price changes are lower in periods of low inflation. And the survey of empirical work on product substitution bias provided by Triplett (1975) shows some evidence that product substitution bias is higher in periods of high inflation.

Outlet Substitution Bias

Outlet bias is driven by the extent of substitution between stores in response to relative price movements and the extent of any relative price movements. One might expect small independents to lower prices more slowly during a recession, because they have less margins to absorb. And one could also expect greater substitution in response to relative price movements during periods of weak activity, because the relative price of searching (i.e. leisure time) has fallen, raising the incentive to search for cheaper options. Consequently outlet bias may be contracyclical.

Quality Bias/ New Goods Bias

It is hard to see many *a priori* reasons for movements in these biases by state. However, it is possible that innovations will be released into the market less frequently during recession, reducing the prospects for new goods bias.

We should stress that this section incorporates qualitative assessment alone. Future work might focus on some of the empirical issues - for example how cyclical is the introduction of new goods?

8. Conclusion

We have estimated a series of 'plausible ranges' for systematic bias in the UK's RPI, where bias is defined as the difference between the actual RPI and a measure of the rate of change of prices for all goods and services currently being consumed. Summing the plausible lower and upper bounds yields an overall range of 0.35 to 0.8pp per annum, excluding the flow of services element of quality bias. Table 10 gives the breakdown alongside the results reached by Crawford and Lebow *et al*.

Table 10

	UK - guesstimate	USA - Lebow <i>et al</i>	Canada - Crawford
Product substitution bias	0.05 to 0.1%	0.2%	0.1 to 0.2%
Outlet substitution bias	0.1 to 0.25%	0.1 to 0.8%	0.08%
Quality adjustment bias	0.2 to 0.3%	0.3%	0.2%
New goods bias	0.0 to 0.15%	0.5%	less than 0.1%
Total	0.35 to 0.8%	up to 1.8%	up to 0.6%*

*Crawford notes that most of the judgements he makes are generous, so that it is likely that total bias is less than 0.5%.

A key point is that the UK and Canadian results are significantly less than those for the US. This probably reflects the assumptions underlying the study quoted, more than real differences - a recent study by the Congressional Budget Office estimated bias of up to 0.8pp per annum in the United States.

Product substitution bias is likely to be lower in the UK than in Canada or the US - given the relative frequency of weighting revisions. However, the differences are fairly small.

The bulk of the differences between the estimates are due to the outlet bias and new goods bias estimates. There are no *a priori* reasons for differences of this scale - they reflect differences in underlying assumption rather than environment.

The main difference between Lebow *et al* and the other estimates lies in the outlet bias estimate. But our simulations suggest that a lower range is robust to a fairly wide ranging set of assumptions - even under extreme assumptions about both market share changes and relative price movements the simulations yielded a figure below Lebow's. *And our estimates may overstate the bias, if relative prices do not continue moving as they did during the late 1980s and early 1990s.* Lebow's estimate of new goods bias is also significantly higher, despite using a technique broadly similar to Crawford's. However, the high figure relies on an (implausibly) high estimate of the proportion of high tech expenditure spent on new goods. Again, the lower estimates appear fairly secure.

Although the three estimates are fairly similar, the quality bias results are the least secure. All three estimates are based on Gordon's work. *But his figures do not allow for the flow of services aspect of durable goods consumption. By ignoring this, both he and we may understate the true bias.* If we extrapolate from Nordhaus, the upper bound for quality bias would be raised from 0.3pp to 0.8pp pa hence adding 0.5pp to the overall upper bound.

While our 'plausible upper bound' is based on simulation and extrapolation, rather than detailed estimation, it is fairly robust to specification changes. More precise estimation requires access to data which is not presently available within the Bank.

Appendix 1 - Simulation Techniques

A1. New Goods Bias

The simulation was designed to assess the impact of various assumptions on the incidence of new goods bias. Both relative prices and the proportion of each sector exposed to the bias were varied. The simulation was based on published monthly data, from 1987 to 1994, both in total and at a sectoral level. New goods bias is assumed to impact on the audio-visual and electrical appliance sectors only. For each, a simulated index was calculated. These indices were substituted for the actual indices and a simulated RPI was calculated. The simulated indices were a weighted average of the actual index (representing established goods) and a new goods index. The new goods weight was initially set at zero, and allowed to grow at a constant rate to reach a terminal level by 1991, when it is assumed that the new good enters the RPI (we experimented with varying the terminal value). Prior to entry into the RPI sample, the new goods index is calculated by reducing prices relative to the all items RPI. In 1991 (when the new good enters) the new goods price index changes at the same rate as the actual index (so that relative prices are assumed unchanged). The overall simulated RPI is calculated by chain-linking the simulated series and published series for other sectors.

A2. Outlet Bias

The simulation was designed to assess the impact of various assumptions on outlet substitution bias - relative price and market share assumptions were varied. The simulations were based on published monthly data for the RPI, both in total and a sectoral level. The foundation of the model is the

observation that the RPI may be viewed as a weighted sum of indices derived from data contributed by different stores types. In this case

$$RPI = \alpha RPI_m + \beta RPI_i + (1 - \alpha - \beta) RPI_d \quad (a1)$$

The shop type indices (RPI_m , RPI_i , and RPI_d for established multiples, independents and new discount stores respectively) are constructs calculated by scaling the published series. The scaling factor was varied according to assumptions about movements in relative prices - note that the model compares endpoints, so there is no need to adjust the scale over time. Comparing an absolute scale of 0.25 yields similar results to these when an annual shift of 3.125% is considered (over 8 years). The weights were also allowed to vary over time, according to assumptions about shifting market share. However, it was necessary to impose a constraint, so that the constructs were consistent with the base series, at the start of the simulation.

$$\beta = 1 - \alpha \quad (a2)$$

The initial weights can now be derived by rearranging equation (a1) in the light of (a2):

$$\alpha = \frac{RPI - RPI_i}{RPI_m - RPI_i} \quad (a3)$$

Assumptions about changes in market share can be incorporated from the initial value derived above. Note that it is changes in the level of market share, rather than its growth rate which are imposed. An overall RPI is derived by chain-linking the derived series and the other constituent series of the RPI, using the published weights.

Appendix 2 - Inputs to the Outlet Bias Simulations

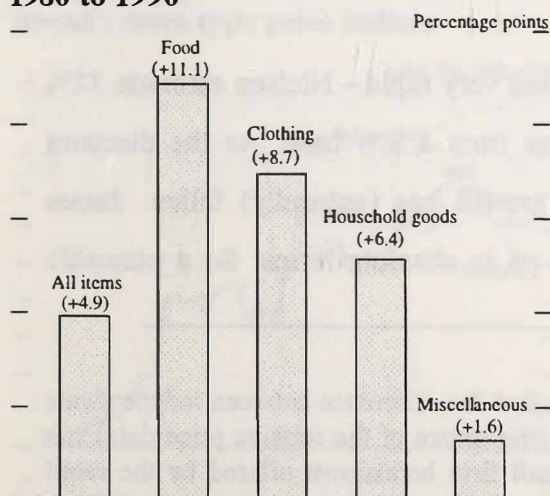
The simulations involve judgements in two areas - the shift in market share between store types and the annual movement in relative prices. Separate judgements may be made for independents relative to established multiples and new discount stores relative to multiples.

Market Share Changes

The past decade has seen a continued expansion of large retailers, at the expense of small independent retailers. Between 1980 and 1990, large retailers market share rose from 67.5% to 78.6% - or by around 1.1% pa. As Chart A.i. shows, the magnitude of this increase in market share was unique to the food sector. However, across all items, large retailers' market share grew by 4.9pp - roughly ½pp per annum.

Chart A.i.

Change in large firms' market share 1980 to 1990*



* Large firm defined as having turnover greater than 2mn in 1984 and 3mn in 1990.

Source : CSO Business monitor

Approximate annual market share changes, derived from chart A.i., are used as the base scenario for market share changes in the various retailing sectors.¹⁶ If we expect the trends of the 1980s to persist, we may expect independent's market share to continue to fall by around ½% pa. Extrapolation from the sizeable (*circa* 1% pa) shift in food retailers' market share is used as an upper bound.

The most simplistic scenario assumes that the RPI sample does not change, so that the entire market share change contributes to bias. As we noted in the main text, sample turnover is 'slow' and also unmeasured.

During the 1980s, supermarkets prospered. With a rapid rise in operating margins, the three large players (Tesco, Sainsbury and Argyll) expanded significantly - the preferred vehicle for growth was an expansion of superstores. As the 80s turned to the 90s, competition began to increase. The industry was dominated by the major retailers. However, a new brand of lower priced competition was attracted by the relatively high margins enjoyed by the food retailing sector in the UK (relative to elsewhere in Europe). The 1990s has seen an increase in the market share of these discount stores.

Initially, growth of the discount sector was very rapid - Nielsen estimate 32% growth during 1992. However, this was from a low base. As the discount sector has become larger, its rate of growth has (naturally) fallen. James Capel's project growth of some 1.25% pa in absolute terms. So a plausible

¹⁶ Note that this analysis implicitly assumes that the difference between independents and multiples in the CSO's RPI breakdown (the source of the relative price data) can be represented adequately by the large/ small firm breakdown offered by the retail sales data.

scenario might show discount stores' market share growing by around 1pp pa. So far, the increase in the discount stores' market share has been largely at the expense of small retailers and independents. Consequently, a plausible scenario would show discount stores' market share growing at the expense of independents.

Relative Price Changes

The CSO has undertaken some research in this area. Separate independent and multiple price indices are available for 2 food items. Point data is also available for a wider range of commodities - for October 1993 and October 1994. These can be used to assess average annual relative price movements. Chart A.ii. plots the separate bread price indices from 1990, to show that multiples prices have been persistently lower. Although they may not persist in the long run, the trend may persist for some time. Chart A.iii. (overleaf) compares annual inflation rates of the bread price indices.

Chart A.ii.

Bread : store type price indices

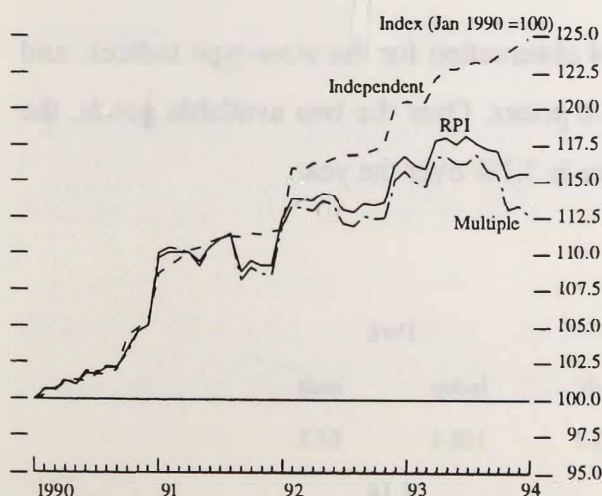
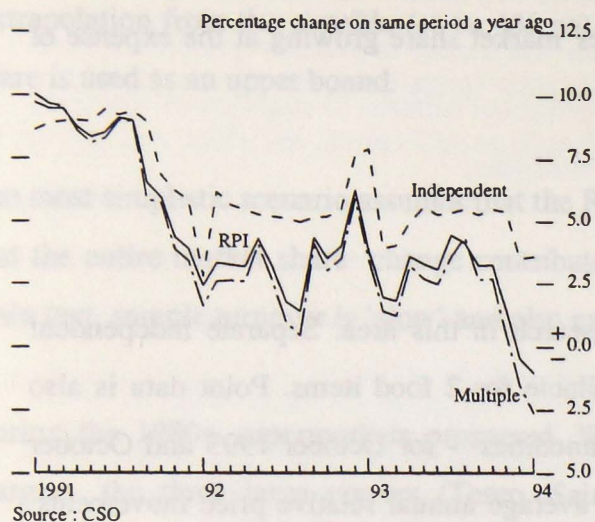


Chart A.iii.

Bread : store type price inflation



Using the CSO's data, we compare the endpoints of the multiple and independent bread price indices. This yields information on the total discrepancy accrued over 4 years, and is used as the basis of a rough analysis of relative price movements. This rough analysis is used to suggest plausible inputs for the scenarios.

Table A.i shows the January 1994 observation for the store-type indices, and the implied annual shift in relative prices. Over the two available goods, the average difference in relative prices is 3.1% over the year.

Table A.i

	Bread		Pork	
	Indep	mult	Indep	mult
Jan 94	124.8	112.7	108.4	94.7
Ind/mult	1.11		1.14	
rel P pa	2.75		3.5	

Data on relative price movements between discount stores and established multiples is relatively scarce. One possibility is to make the arbitrary assumption that the difference in relative prices between independents and multiples is replicated between discount stores and mainstream multiples. Given the *raison d'être* of the discount store (pricing lower than others), this assumption may appear to understate relative price differentials. However, there are two arguments in favour of its adoption. First, outlet bias revolves around quality-adjusted price differentials. Since discount stores are widely perceived to be of lower quality, the quality adjusted price differential will be lower than at first sight. Second, the mainstream retailers' response to discount stores has been to boost sale of own brands and cut prices, thus reducing any discrepancy in relative price changes.

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