


Prices and Inequality: The United Kingdom Experience

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PRICES AND INEQUALITY: THE UNITED KINGDOM EXPERIENCE[1]

I. Introduction

This paper shows that relative consumer price changes in the United Kingdom since 1964 have had an inequality-increasing bias. This is in line with findings by earlier investigators. However, this paper uses different methods in establishing this result and embodies a number of conceptual and empirical innovations.

The realisation that if the prices of " luxuries " and " necessities " move differently, then different groups are differently affected is, of course, not new. On the recommendation in 1968 of the Cost of Living Advisory Committee, the Department of Employment publishes quarterly in its Gazette one- and two-person pensioner price indices (excluding housing) going back to 1962, as well as the General Index of Retail Prices. These show that the cost of living of pensioners has been increasing more rapidly than the Retail Price Index. Some specific studies of the cost of living for different income groups have also been made. Lydall (1959) found that over the period 1938–49–57, high income households (the top 3·6%) had a smaller rise in the cost of living than other households excluding pensioners. However, Seers (1951) and Allen (1957, 1958), with somewhat different data, showed the opposite tendency for the early part of the period. For the period 1951–56, Brittain (1960) found that the lower the income group, the more unfavourable the price trend so that the egalitarian war-time price trends were almost completely wiped out. Lynes (1962) found that the cost of living for his sample of poor fatherless families increased very substantially more than the Index of Retail Prices over the period 1948 to 1961. Tipping (1970) for 1956–66 found that a household at the 5% percentile from the bottom experienced 6% more inflation than one at the 95% percentile and 4·1% more than one at the 25% percentile. However, one serious problem with that study (and some of the earlier ones) is that household size is ignored: a given income group includes both relatively badly off large households and relatively well off small households.

One important advantage of the present paper is that household composition is explicitly taken into account in accordance with the theory developed in Muellbauer (forthcoming). Others lie in the use of a system of demand equations. This enables the fixed weight price indices used in all

[1] I am grateful for invaluable computational assistance to Bernard Pearson and Peter Okell, to Angus Deaton for providing parameter estimates without which the paper could not have been written, to Bertie Hines, Tony Atkinson, Carol Nussey, Ben Fine and seminar groups at Warwick University and the London School of Economics for valuable comments. I remain responsible for any remaining misinterpretations and ambiguities.

the above studies to be superseded by " true " (*i.e.*, constant utility) cost of living indices which permit consumers to substitute in response to relative price changes. Further, cost of living indices corresponding to EVERY income (or rather expenditure) level can now be constructed. In the fixed weight approach the expenditure shares for each income level would have to be observed. In practice, these are subject to much sampling variability and all the cited studies distinguish only very few broad income categories.[1] None of them has been able to arrive at an overall evaluation based on a comparison of the money and the real income distribution of the importance of the bias in relative price changes. The present paper, however, does attempt to do this. Following the procedures put forward in Muellbauer (1973), the money and the real (in 1964 prices) distribution of household expenditure for 1970 is compared with that for 1964. Part of the reason for selecting these dates is to attempt to throw some light on the controversy[2] over what happened to inequality under the 1964–70 Labour government.

At first sight the evidence suggests that the bias in relative price change was not sufficient to reverse the small but significant reduction in money inequality over 1964–70 which my data reveal. However, as I shall show, the biases in my measurements are almost all in the direction of *understating* the inegalitarian bias in relative price changes. Moreover there are, I argue, measurement errors in the Family Expenditure Survey (FES) which understate relative inequality in 1970. Thus, whether inequality increased or decreased over 1964–70 is still an open question.

The paper is divided into 6 sections. Section II discusses the data for 1964 and 1970 of the money expenditure distribution corrected for household size and presents the data on prices which underly the empirical results. Section III briefly discusses those aspects of the theory of true cost of living and real income indices and household composition effects which are relevant to an appreciation of the contents of this paper. Section IV presents estimates of the linear expenditure system (LES) of demand equations and the implied true cost of living indices for different expenditure levels. Section V presents the money expenditure distributions corrected for price changes and attempts to quantify the price change component in the change in inequality. Conclusions are presented in section VI. In the first appendix, I discuss various types of possible biases in the results and attempt to evaluate the size and direction of some of them. Section A treats some difficulties with the LES, in particular the possibility of some quantities demanded becoming negative, and with the durables equation, and those associated with household composition effects. Section B discusses problems caused by the

[1] Though Tipping (1970) uses an interpolation technique which can be used to obtain a finer discrimination over incomes.

[2] See the review by Townsend of Stewart's (1972) position published in Beckerman (1972) in *The Listener*, April 27th, 1972, and the succeeding correspondence in the letter pages of *The Listener* until July 27th, 1972.

breadth of the commodity groups for which the LES was estimated. Two difficulties in the Family Expenditure Survey are discussed in section C. These are the treatment of housing in the FES and the problems this poses for the development of meaningful inequality measures and the under-representation of the higher and the lowest income groups in the FES in 1970 compared with 1964. Appendix 2 investigates the effects of capital gains in housing and suggests that those of 1970–72 were enough to overshadow the entire measured reduction in inequality of 1964–70.

II. The Data

(a) *Prices*

The price data, with sources and definitions, from which in Section 4 I calculate constant utility cost of living indices corresponding to different expenditure levels, are presented in Table I. These price indices are

TABLE I

Paasche Price Indices for 1963–1972

	(1) Food.	(2) Cloth-ing.	(3) Hous-ing.	(4) Fuel and Light.	(5) Drink and Tobacco.	(6) Travel.	(7) Misc. Goods.	(8) Misc. Ser-vices.	(9) Dur-ables.	Total Exp.
1963	1·000	1·000	1·000	1·000	1·000	1·000	1·000	1·000	1·000	1·000
1964	1·026	1·015	1·067	1·028	1·055	1·028	1·035	1·028	1·012	1·033
1965	1·062	1·036	1·138	1·050	1·155	1·089	1·073	1·072	1·030	1·080
1966	1·096	1·063	1·211	1·087	1·190	1·130	1·107	1·135	1·042	1·122
1967	1·117	1·080	1·253	1·101	1·206	1·170	1·127	1·190	1·063	1·150
1968	1·151	1·096	1·300	1·164	1·248	1·253	1·240	1·257	1·114	1·202
1969	1·214	1·138	1·369	1·162	1·343	1·322	1·285	1·333	1·156	1·267
1970	1·274	1·200	1·469	1·170	1·394	1·380	1·386	1·417	1·234	1·335
1971	1·396	1·284	1·605	1·230	1·453	1·497	1·512	1·526	1·321	1·439
1972	1·491	1·375	1·740	1·303	1·507	1·522	1·547	1·626	1·371	1·516

Source: 1963–71: from National Income and Expenditure Blue Book, 1972, Tables 22, 23. Implicit money expenditure deflators.

Columns 3, 4, 9: as in the tables.

Columns 1, 2: as in the tables but with income in kind allocated in ratio 5·0 : 2·3.

Column 5: sum of alcoholic drink and tobacco.

Column 6: running costs of motor vehicles, travel, communications services, consumer expenditure abroad.

Column 7: other household goods and books, newspapers and magazines, chemists goods, misc. recreational goods and other misc. goods.

Column 8: entertainment and recreational services, domestic service, catering, wages in non-profit making bodies, insurance and other services.

1972 prices indices from Monthly Digest of Statistics, March 1973 and provisional figures made available privately by the CSO for columns 6, 7, 8.

basically Paasche indices obtained by dividing, for a nine commodity group breakdown, the money expenditures by real expenditures in 1963 prices. Most other investigators have not used this Blue Book data but instead have worked with the components of the Index of Retail Prices.[1]

[1] The Index of Retail Prices shows more inflation than the consumer expenditure deflator. This may be in part a bias caused, it has been argued, by some inflexibility in the system of collecting

It is interesting to have a preliminary look at whether prices of " necessities " have indeed increased more than those of " luxuries." The last column of Table II (on page 40) gives total expenditure elasticities for 1964 mean total expenditure computed from the estimated demand equations which are discussed in Section IV. It is clear that housing and miscellaneous services are " necessities " (total expenditure elasticities less than unity) and have increased in price more over 1963–72 than the average, and durables are " luxuries " and have increased in price less than the average. Over 1970–72 the relationship is even stronger[1] with the large increase in food prices. However, these figures alone are not totally convincing on the inegalitarian nature of price changes. To examine this properly it is necessary to construct cost of living indices for different expenditure levels.

(b) *Expenditure Distributions*

To evaluate the overall significance of relative price movements on inequality, it is necessary to have some distribution data. I shall examine the money expenditure distributions for 1964 and 1970 using FES data. It is total money expenditure rather than money income which is the relevant concept for the non-intertemporal framework[2] adopted in this paper.

Unfortunately, detailed expenditure distribution data by household type were not directly available for both 1964 and 1970. Instead I took the tables[3] of the respective income distributions by household size as my starting point. To these, I applied estimated expenditure/income ratios obtained from the detailed expenditure by household type tables in 1964 and 1970.[4] Combining the two sets of data, the money expenditure distribution for each household size category is obtained.[5]

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price information. It may take insufficient account of consumers' shifting to new and cheaper lines and retail outlets. The consumer expenditure deflator is more of a unit value index and hence is less affected by this.

[1] The Spearman rank correlation coefficient for the ordering of the total expenditure elasticities and the 1963–72 price changes is — 0·4 which is not significant at the 0·05 level. For 1970–72 price changes the correlation coefficient is — 0·61 which is significant. However, such simple statistical criteria give only a superficial indication of the bias in relative price change.

[2] This is not to deny that savings behaviour may be affected by relative price changes, say in housing, nor to say that some intertemporal index number concepts cannot be defined. However, the burdens both on the additional data requirements and of the additional assumptions that would be necessary, are too heavy to contemplate at this stage.

[3] Table G, p. 7 and Table 33, p. 94 from the Annual Reports of the FES in 1964 and 1970 respectively.

[4] Tables 8–12, pp. 70–86 and Tables 4–8, pp. 30–50 for 1964 and 1970 respectively.

[5] The main difficulty with this method is that while in 1964 household types were given in categories of numbers of persons per household, in 1970 the household size categories were instead presented in the following categories: (*a*) 1 adult; (*b*) 1 man, 1 woman; (*c*) 1 man, 1 woman, 1 child; (*d*) 1 man, 1 woman, 2 children; (*e*) 1 man, 1 woman, 3 or more children. However, the household size categories for 1964 are predominantly composed of the same types of families. For example, a four person family has 1·86 children under 16 on average and a two person household has 0·06 children. Thus it is reasonable to assume that similar relationships between income categories and average expenditure apply for both kinds of household size categories.

The next step is to pool these data into a single household expenditure distribution. This involves adjusting the relative expenditures of households of different sizes to bring their welfare levels onto the same monetary yardstick. In the next section I summarise the method for doing this put forward in Muellbauer (forthcoming). Under the particular simplification which I adopt in the current paper, the familiar technique of deflating household expenditure by " adult equivalent scales " is consistent with this theory. In fact, I have chosen the scales used by Prest and Stark (1967) and Stark (1972). They are:

Size of unit in persons	1	2	3	4	5	6	7	8
Equivalence scale	1	1·6	2·1	2·5	2·8	3·2	3·6	4·0

This implies, for example, that a two person household with £1,600 is as well off as a one person household with £1,000. These scales are an average of 1964 National Assistance Board scales and estimates by Jackson and Nicholson—see Stark (1972), pp. 51–3 for details. They are not necessarily optimal[1] but appear sensible. For measuring *changes* in inequality the method is probably quite robust.

III. Some Index Number Concepts

Let $y = m(p, u)$ be as in Muellbauer (1973), the outlay required by an individual to reach utility level u at the price vector p. $m(.)$ is known as the expenditure function. Let $u = V\left(\frac{y}{p}\right)$ be the indirect utility function, where $\frac{y}{p} = \left(\frac{y}{p_1}, \ldots, \frac{y}{p_r}\right)$. This gives attainable utility in terms of total outlay and prices and can be found by substituting the Marshallian demand functions into the direct utility function $u = U(q_1, \ldots, q_r)$. A true (constant utility) cost of living index which compares prices p_1 and p_o is given by

$$\frac{m(p_1, u)}{m(p_o, u)} \quad . \quad . \quad . \quad . \quad . \quad (1)$$

There is more than one sensible level of u which can be chosen. I have chosen to work with the base period concept $u_o = V\left(\frac{y_o}{p_o}\right)$. A real expenditure index compares the expenditure necessary to purchase two utility levels u_o and u_1 at a given reference price vector p and is given by

$$\frac{m(p, u_1)}{m(p, u_o)} \quad . \quad . \quad . \quad . \quad . \quad (2)$$

In Muellbauer (forthcoming), I have shown how these concepts may be extended to comparisons between households of different compositions.

[1] In principle, they could be estimated by, say, maximum likelihood methods from FES data.

Briefly, this is done by introducing parameters $(m_1, \ldots, m_r)$ called specific household equivalent scales which play a role analogous to prices. Let the utility function for household H be

$$u_H = U\left(\frac{q_1}{m_{1H}}, \ldots, \frac{q_r}{m_{rH}}\right) \quad . \quad . \quad . \quad (3)$$

Then the expenditure function is given by

$$y_H = m(p_H^*, u_H) \quad . \quad . \quad . \quad . \quad (4)$$

and the indirect utility function is given by

$$u_H = V\left(\frac{y_H}{p_H^*}\right) \quad . \quad . \quad . \quad . \quad (5)$$

where

$$p_H^* = (p_1 m_{1H}, \ldots, p_r m_{rH})$$

Relative "real expenditure" levels of households H and J are then given by

$$\frac{m(p_H^*, u_H)}{m(p_H^*, u_J)} \quad \text{or} \quad \frac{m(p_J^*, u_H)}{m(p_J^*, u_J)} \quad . \quad . \quad . \quad (6)$$

depending on which household's "price vector" is taken as reference. Typically a one adult household would be taken as reference. These indices can therefore be used to convert a money household expenditure distribution to a real per adult equivalent expenditure distribution.

For the purpose of this paper I have made the assumption (which is empirically testable) that

$$\frac{m_{iH}}{m_{iJ}} = \frac{m_{oH}}{m_{oJ}}, \quad i = 1, \ldots, r \quad . \quad . \quad . \quad (7)$$

This means that economies of scale in households affect all goods proportionately. It implies that (4) can be written

$$y_H = m_{oH} \cdot m(p, u_H) \quad . \quad . \quad . \quad . \quad (8)$$

Thus, if two households H, J have the same utility, then

$$\frac{y_H}{m_{oH}} = \frac{y_J}{m_{oJ}} \cdot \quad . \quad . \quad . \quad . \quad (9)$$

It should be obvious that m_{oH} and m_{oJ} have precisely the same meaning as the adult equivalent scales used in Section 2. Thus if H is a two person, J a three person household, they have the same welfare level if

$$\frac{y_H}{y_J} = \frac{m_{oH}}{m_{oJ}} = \frac{1{\cdot}6}{2{\cdot}1}$$

What is more, this result follows at any price vector p, which would not in general be true of (6). Condition (7) is the only one under which the con-

version to adult equivalents can proceed independent of prices. Once we have carried out this conversion, we can treat each adult equivalent unit as if it were a separate individual and use the indices (1) and (2) for cost-of-living and real income indices.

IV. The LES and True Cost-Of-Living Indices for Various Total Expenditure Levels

It is assumed that for a one adult household, the utility function is given by

$$u = \prod_{i=1}^{r} (q_i - \alpha_i)^{\beta_i}, \quad \Sigma \beta_i = 1 \; . \qquad . \qquad . \qquad (10)$$

with corresponding demand functions[1]

$$q_i p_i = \alpha_i p_i + \beta_i (y - \Sigma \alpha_j p_j) \quad . \qquad . \qquad . \qquad (11)$$

Then for some other household H, it follows[2] that

$$u_H = \prod_{i=1}^{r} \left(\frac{q_{iH}}{m_{oH}} - \alpha_i \right)^{\beta_i}, \quad \Sigma \beta_i = 1 \qquad . \qquad . \qquad (12)$$

The corresponding demand equations are

$$q_{iH} p_i = \alpha_i m_{oH} p_i + \beta_i \left(y_H - \sum_{j=1}^{r} \alpha_j m_{oH} p_j \right) \quad . \qquad . \qquad (13)$$

If both quantities and incomes are divided by the adult equivalent scale m_{oH}, the demand equations have the form

$$\frac{q_{iH} p_i}{m_{oH}} = \alpha_i p_i + \beta_i \left(\frac{y_H}{m_{oH}} - \sum_{j=1}^{r} \alpha_j p_j \right) \qquad . \qquad . \qquad (14)$$

Thus assuming different households have the same values (α_i, β_i), face the same prices and have incomes large enough so that $q_i \geqslant 0$, all i, then the (α_i, β_i) can be estimated without aggregation error from aggregate data if the quantity expenditure data are deflated by the adult equivalent population. The utility function (10) thus corresponds to an adult equivalent unit.

Substituting the demand functions (11) into (10) we obtain the indirect utility function

$$u = V\left(\frac{y}{p}\right) = (y - \Sigma \alpha_i p_i) \; \Pi \left(\frac{\beta_i}{p_i} \right)^{\beta_i} \qquad . \qquad . \qquad (15)$$

and the expenditure function

$$y = m(p, u) = \Sigma \alpha_i p_i + u \; \Pi \left(\frac{p_i}{\beta_i} \right)^{\beta_i} \qquad . \qquad . \qquad (16)$$

[1] See Stone (1954) for discussion of these functions.

[2] This is so from (7) where m_{oJ}, m_{iJ} are normalised at unity for the adult reference household.

The parameters have attractive interpretations: for good i, α_i is the " committed purchase " and β_i is the marginal propensity to consume out of total expenditure. " Necessary " goods are those with low expenditure elasticities, *i.e.*, high α_i and low β_i. " Luxuries " have low α_i and high β_i. This is extremely helpful in interpreting the implied cost-of-living indices.

Let $a_t = \Sigma\, \alpha_j p_{jt}$ and $b_t = \Pi \left(\frac{p_{it}}{\beta_i}\right)^{\beta_i}$. Then (15) and (16) become respectively

$$u = (y - a)b^{-1} \quad . \quad . \quad . \quad . \quad (17)$$
$$y = a + ub \quad . \quad . \quad . \quad . \quad (18)$$

Thus the true cost-of-living index (1) becomes

$$\frac{m(p_t, u_{so})}{m(p_o, u_{so})} = \frac{1}{y_{so}}\left[a_t + (y_{so} - a_o)\frac{b_t}{b_o}\right] \quad . \quad . \quad (19)$$

where s refers to the s-th adult or adult equivalent unit. The RHS of (19) can be rewritten as

$$\left(\frac{a_o}{y_{so}}\right)\frac{a_t}{a_o} + \left(1 - \frac{a_o}{y_{so}}\right)\frac{b_t}{b_o} \quad . \quad . \quad . \quad (20)$$

Thus the cost-of-living index is a weighted average of $\frac{a_t}{a_o}$ and $\frac{b_t}{b_o}$. But $\frac{a_t}{a_o} = \frac{\Sigma\, \alpha_i p_{it}}{\Sigma\, \alpha_i p_{io}}$ is an arithmetic price index using " committed purchases " as weights. Thus it gives " necessities " high weights. And $\frac{b_t}{b_o} = \Pi \left(\frac{p_{it}}{p_{io}}\right)^{\beta_i}$ is a geometric price index using marginal propensities to consume as weights. Thus it gives high weights to " luxuries." From (20) it is obvious that $\frac{b_t}{b_o}$ tends to dominate for a rich person and $\frac{a_t}{a_o}$ tends to dominate for a poor person.[1]

We now turn to the estimates. The demand functions estimated by Angus Deaton have time trends in the β_i's, *i.e.*, $\beta_i = \delta_i + \text{time} \times \gamma_i$. The 9 commodity breakdown is as defined in Table I. His data came from the 1971 Blue Book and consisted of annual real expenditures in 1963 prices for 1954–70 deflated by mid-year population.[2] The estimating technique used was identical to that in Parks (1969). It is maximum likelihood applied to a general contemporaneous covariance matrix of disturbances and takes into account the singularity of this matrix. No serial dependence is allowed for.

[1] Note that there is no fundamental difficulty if $a_o > y_{so}$. Then a_t/a_o has a weight exceeding unity and b_t/b_o has a negative weight and the " committed expenditure " interpretation of a_o is no longer valid. But this does not matter. As long as $q_i \geqslant 0$, all i, the expenditure function satisfies all the fundamental conditions of consumer theory (concave in p, increasing in p and u) and hence the corresponding cost-of-living indices are also valid.

[2] It would have been more appropriate to deflate by the number of adult equivalents but since there was very little structural change in the population, this does not cause problems.

The individual equations fit very well, with R^2 usually in excess of 0·98.[1] The estimates are presented in Table II.

TABLE II

Linear Expenditure System Estimated by Angus Deaton of the Department of Applied Economics, Cambridge

$$q_i p_i = \alpha_i p_i + \beta_i \left(y - \sum_{j=1}^{9} \alpha_j p_j \right),$$

$$\sum_{1}^{9} \beta_i = 1 \quad \text{and} \quad \beta_i = \delta_i + \text{time} \times \gamma_i$$

	α_i.	δ_i.	γ_i.	Total expenditure elasticity.*
1. Food	76·36	0·1090	−0·0004	0·47
2. Clothing	21·26	0·1258	−0·0003	1·33
3. Housing	34·23	0·0530	0·0006	0·50
4. Fuel and light	12·74	0·0442	0·0002	0·94
5. Drink and tobacco	34·95	0·1040	−0·0001	0·83
6. Travel	21·45	0·1178	0·0006	1·27
7. Misc. goods	21·03	0·1271	−0·0002	1·33
8. Misc. services	39·76	0·1126	0·0004	0·82
9. Durables	6·51	0·2064	−0·0008	2·46

* Total expenditure elasticities computed for 1964 at mean total expenditure in 1964.

Data from Blue Books, categories as in Table I, 1954–70 in 1963 prices. Expenditures deflated by mid-year population.

In 1964, durable expenditure becomes zero at a level of expenditure = £5·97/adult equivalent/week. About 16% of the adult equivalent population had expenditure less than £5·97. Expenditure on any other category becomes zero at levels of total expenditure experienced only by the poorest 1 or 2%.

In Table III, values of $\frac{a_t}{a_o}$, $\frac{b_t}{b_o}$ and cost-of-living indices corresponding to various base year expenditure levels are presented.[2] It is quite clear that relative price changes from 1964 to 1972 have persistently favoured the better-off. The 1971–72 change is particularly marked in this respect.

One feature of Table III needs to be discussed. As we shall see in Appendix 1A below, these parameter estimates imply that durable purchases became zero in 1964 at a weekly adult equivalent total expenditure of £5·97. It is necessary to assume that different preference patterns and demand functions, defined over the remaining 8 goods, apply for these consumers. Assuming that the same parameter values hold, we have to redefine the marginal propensities to consume:

$$\beta_i^* \equiv \frac{\beta_i}{\sum_{j=1}^{8} \beta_j}, \quad \text{so that} \quad \sum_{i=1}^{8} \beta_i^* = 1$$

[1] Further details will be available in a forthcoming monograph by Angus Deaton.

[2] Since the *mpc*'s were estimated with time trends, the values corresponding to some *given* year must be chosen so that the indices are ordinal. I chose 1964, but, in the event, choosing 1970 gave virtually the same results.

TABLE III

$\frac{a_t}{a_o}$, $\frac{b_t}{b_o}$ and Cost-of-Living Indices for Different Expenditure Levels

	Over 9 goods.		Over 8 goods† (excl. durables).		Cost-of-living indices for different 1964 weekly expenditure levels.								
					1964 Expenditures in £								
	$\frac{\Sigma \alpha_i p_{it}}{\Sigma \alpha_i p_{io}}$	$\Pi\left(\frac{p_{it}}{p_{io}}\right)^{\beta_i}$	$\frac{\Sigma \alpha_i p_{it}}{\Sigma \alpha_i p_{io}}$	$\Pi\left(\frac{p_{it}}{p_{io}}\right)^{\beta_i *}$	2·39	3·19	4·79	6·39	9·56	14·37	19·16	38·32	57·47
1964	1·000	1·000	1·000	1·000	1·000	1·000	1·000	1·000	1·000	1·000	1·000	1·000	1·000
65	1·048	1·041	1·049	1·047	1·053	1·051	1·050	1·048	1·046	1·044	1·043	1·042	1·042
66	1·090	1·074	1·091	1·086	1·100	1·096	1·093	1·090	1·085	1·081	1·080	1·077	1·076
67	1·118	1·100	1·119	1·113	1·129	1·125	1·121	1·118	1·112	1·111	1·106	1·103	1·102
68	1·168	1·156	1·170	1·170	1·169	1·169	1·169	1·168	1·164	1·161	1·160	1·158	1·157
69	1·229	1·210	1·232	1·228	1·238	1·235	1·233	1·230	1·223	1·219	1·217	1·213	1·212
70	1·296	1·278	1·298	1·294	1·304	1·302	1·299	1·296	1·290	1·286	1·284	1·281	1·280
71	1·399	1·375	1·401	1·393	1·415	1·411	1·404	1·400	1·391	1·386	1·383	1·379	1·378
72	1·481	1·443	1·484	1·466	1·514	1·502	1·490	1·482	1·469	1·460	1·456	1·449	1·447
Ratio expenditure/mean expenditure					0·25	0·33	0·50	0·67	1·00	1·5	2·00	4·00	6·00
Approximate % position in distribution					less than 5%		5–10%	18·4%	67·1%	87·8%	93–98%	98–100%	

† β_i^* is defined as $\frac{\beta_i}{\sum_{i=1}^{8} \beta_i}$

Durables demand becomes zero for expenditure less than £5·97/week.

The α's are unchanged. The cost-of-living indices for base year total expenditure below £5·97 in Table III are defined correspondingly. The values for $\frac{a_t}{a_o}$ and $\frac{b_t}{b_o}$ defined for 8 goods reveal more inflation. This is as expected since durables had below average price increases.[1]

V. Inequality of Real Expenditures and the Price Component in Inequality Change

Following the analysis in Section III, the 1970 money expenditure distribution for adult equivalents was converted into 1964 prices by means of the following real expenditure indices:

$$\text{real expenditure for adult equivalent unit } s = m(p_o, u_{st}) = a_o + u_{st} \cdot b_o$$

where

$$u_{st} = (y_{st} - a_t)b_t^{-1}$$

o refers to 1964, t to 1970.

To evaluate the quantitative significance of the relative price changes, I calculated inequality indices for the money and the real expenditure distributions. Following the approach of Atkinson (1970a, b), let us consider a strictly quasi-concave, symmetric social welfare function defined on expenditures. Then an intuitively sensible approach to measuring inequality of a given total is to see how far " actual social welfare " lies below " optimal social welfare " (*i.e.*, social welfare for an equal distribution). Atkinson's index I is based on this notion. Assuming that the function is additive[2] and homothetic[3] as well as symmetric and strictly concave, the form is

$$W = \sum_{j=1}^{n} \left(a + \frac{b}{1+\delta} y_j^{1+\delta} \right) . f(y_j)$$

where $f(y_j)$ is the relative frequency of the j-th expenditure group. Then Atkinson's index is:

$$I = 1 - \left[\sum_{j=1}^{n} \left(\frac{y_j}{\bar{y}} \right)^{1+\delta} f(y_j) \right]^{1/1+\delta}$$

[1] This approach is not perfectly satisfactory, however. The reason is that relative price changes affect the utility level at which durable purchases become zero, *i.e.*, the switch-over point between the two preference patterns. Since only a narrow utility band is affected and there is very little difference within it between the two implied cost-of-living indices, I have chosen to ignore this particular problem here and subsequently. However, Appendix 1A discusses the general issues which are involved.

[2] Additivity means that the marginal social rate of substitution between two persons' expenditures is independent of all other persons' expenditures.

[3] Homotheticity means that giving everyone a fixed percentage more leaves the measure unaffected. This specifically contradicts the notion of some subsistence or poverty level of expenditure which one might wish to take as the zero mark of the expenditure yardstick which is relevant for social welfare.

$I = 0$ represents perfect equality. An important attraction of the index is that by varying δ the degree of inequality aversion is varied. For example, $\delta = -2$ implies that if Mr. A has twice Mr. B's income, then the additional social welfare of £1 extra for Mr. B is four times that of £1 extra for Mr. A. If $\delta = -1$, the factor is two instead of four.

I have taken a range of values from $\delta = -0{\cdot}5$ to $\delta = -2{\cdot}5$. Table IV (*b*) presents values of the inequality index for different values of δ for 1964 and 1970 in money and real terms. It shows that in this period real inequality fell somewhat but that the fall in money inequality overstates this reduction by 13–15%. This result is insensitive to the precise value of δ selected.

One interesting feature of the normative approach to inequality measurement is that there is an equivalence between giving every one a certain proportion more real purchasing power (hence keeping inequality constant) and redistributing it less unequally. For example, the estimated reduction in inequality of 0·0151 for $\delta = -2$ is equivalent to giving everyone 1·51% more real income. However, I do not want to make too much of this equivalence—in part because of the objections in footnote 3, p. 42.

No single summary measure of inequality is totally satisfactory. Therefore it is worth looking in more detail at the distribution changes. Quintile shares and their changes are presented in Table IV (*a*). They suggest that

TABLE IV (*a*)

*Quintile Shares of the 1964, 1970 Distributions of Money and Real Expenditure per Adult Equivalent**

Quintile.	(1) 1964 Money exp. share. (%)	(2) 1970 Money exp. share. (%)	(3) 1970 Real exp. share. (%)	(4) (3) − (2) (%)
1st	10·81	11·97	11·90	−0·07
2nd	15·72	15·18	15·14	−0·04
3rd	18·04	18·27	18·25	−0·02
4th	21·68	21·99	22·01	+0·02
5th	33·75	32·58	32·69	+0·11

* The quintile shares are derived from logarithmic interpolation on the cumulative expenditure and frequency distributions. Figures in columns (1) to (3) are less accurate than suggested by the numbers of significant places. Column (4), however, is relatively accurately determined given that columns (1) to (3) are roughly correct.

while the top and bottom quintile respectively decreased and increased their shares, the fourth and second quintiles moved in the opposite, *i.e.*, inequality increasing, direction. The effects of relative price changes on quintile shares are measured by comparing the money and real quintile shares. These effects seem fairly small. It seems likely that changes *within* the top and bottom quintiles, which are ignored here, may be rather important. To that extent the above inequality indices do have advantages.

Incidentally, as an indication of the *absolute* level of inequality, Table IV is misleading. There is a systematic bias in the response rate in the FES so that higher income groups and to some extent the very poorest are under-represented. Thus, Table IV understates inequality.

TABLE IV (*b*)

Inequality Indices for Various Values of δ

δ.	(1) 1964 Money.	(2) 1970 Money.	(3) 1970 Real.	(4) ((3) − (2))	(5) (4)/(1) − (3) (%)
−2·5	0·1676	0·1472	0·1495	0·0023	13
−2·0	0·1387	0·1216	0·1236	0·0020	13
−1·5	0·1080	0·0947	0·0962	0·0015	13
−1·0	0·0749	0·0658	0·0669	0·0011	14
−0·5	0·0391	0·0344	0·0350	0·0006	15

VI. CONCLUSIONS

This paper examines the differential impact on United Kingdom households of price changes in 1964–72. Constant utility cost-of-living indices are calculated for adult equivalent units with different 1964 expenditure levels. These show that the cost of living of the less well off increased more rapidly. The parameters for these indices are obtained from estimates by Deaton of the Linear Expenditure System of demand equations. The price data come from Blue Book sources. Previous studies are based on fixed weight indices which do not permit consumers to substitute in response to relative price changes and primarily use data from the Family Expenditure Survey. However, my conclusions are consistent with previous findings. The one- and two-pensioner price indices published by the Department of Employment show larger increases since 1962 for pensioners than the general public. Lydall (1959) for 1949–57, Brittain (1960) for 1951–56, Lynes (1962) for 1948–61, Tipping (1970) for 1956–66 all find similar inegalitarian trends. My data suggest that since 1970 this tendency has accelerated: Table III shows that the percentage spread between the cost-of-living indices for different expenditure levels widened almost twice as much over 1970–72 as over 1964–70.

No United Kingdom study has, as far as I know, evaluated the consequences of these tendencies for overall measures of inequality and hence measured the price change component in inequality change. This paper attempts to do so. The 1970 money household expenditure distribution is converted to 1964 prices and compared with the 1964 distribution. Household equivalence scales, which have a consistent micro-economic foundation as explained in Muellbauer (forthcoming), are used to bring the incomes of households of different sizes on to a common yardstick. The conclusion is reached that if various measurement errors are ignored, the 1964–70 re-

duction in money inequality overstates the real reduction by 13–15%. This conclusion is not sensitive to the amount of inequality aversion built into the inequality measure which is the same as that suggested by Atkinson (1970).

The whole of Appendix 1 is devoted to examining the effects of various biases and measurement errors. These act almost all to *understate* the inegalitarian tendencies I have found.

My investigations into non-response rates in the FES suggest that the extremes of the distribution were better represented in 1964 and hence give another reason for regarding part of the 1964–70 decline in inequality as spurious. Taking this into account raises the price component in inequality change to about 14–17%.

In addition, there are some effects which I have not been able to quantify.[1] The treatment of housing in the FES imputes rents to owner-occupiers but ignores capital gains. I have also argued that official rent indices understate hardship in the rental market. Both effects cause the inegalitarian bias in relative price changes to be understated. Another is due to the breadth of commodity groups. By far the most serious problem arises in transport which is treated as one good. The fact that public transport has increased much more in price than private transport has a strong inegalitarian effect which, as in Tipping (1970), is here ignored. For reasons explained in Appendix 1A connected with the estimation of the demand equations, it seems likely that the weight of durable goods in the indices for the well to do is underestimated. Since durables have had below average price increases, cost-of-living indices for the well to do are overestimated.

Remembering that in 1964–70 the inegalitarian bias in price changes was small compared to 1970–72, there can be no question about the importance of making inequality comparisons in real rather than money terms.

There are some further considerations which strengthen these conclusions. It seems likely that if insufficient account in official price indices is taken of quality change, the bias tends to understate inequality. Some American studies (see Griliches, 1971) indicate that for many goods whose quality improves, the official price indices overstate the true price increases. These goods tend to be durables and other more processed goods which are consumed in higher proportions by the more well to do. On the other hand there can be little doubt that generally the quality of public transport, for example in frequency of service, has declined.[2] It seems likely that little account of this has been taken in official price statistics. Yet another pointer in this direction is the massive post-1950s public and private investment complementary to the private motor car.

My study confirms that for more than twenty years relative consumer

[1] But see Appendix 2.

[2] An important general point arises here. The allocative mechanism may operate not only through prices but also through queues, " pull," other administrative devices or sheer accidental availability. Prices alone then do not convey anything like all the relevant information. This is even more serious, probably, for housing.

price changes have had an inegalitarian bias and suggests that the degree of bias has recently been increasing. What of the future? Sound economic and political reasons exist for expecting these trends to continue. Increasing attention is being paid to policies to relieve some of the hardship that is involved. It is highly relevant, therefore, to discover which are the really strategic commodities. To this end I have examined the consequences of increasing the relative price of each of the nine commodity groups in turn by 25%.

Table V shows the effect on I, the index of inequality suggested by Atkinson (1970), for various values of the degree of inequality aversion (δ ranges from $-2{\cdot}5$ to $-0{\cdot}5$). I is measured in real terms, *i.e.*, in 1964 prices. ΔI represents the change in real I and ΔI^* is the ratio of this to the 1964–70 estimated change (uncorrected for various measurement errors) to give some idea of the order of magnitudes involved.

Food turns out to be *the* strategic commodity: a 25% rise in the relative price of food would wipe out about half of the measured (uncorrected) decline in real inequality which my data suggest occurred between 1964 and 1970.

Housing is next; it is just over half as important as food. A 25% relative rise would wipe out about 25% of the 1964–70 inequality change. Apart from increased costs of housing the capital gains of house-owners have inegalitarian effects, which, it is suggested in Appendix 2 are overwhelmingly more important. The third most important group is miscellaneous services which includes entertainment and recreational services, insurance and catering. At the other end, a 25% relative rise in durables prices would *reduce* inequality by about 37% of the 1964–70 change. It is important to note that these results are insensitive over a wide range to the degree of inequality aversion built into the inequality index. They are all underestimates because the 1964–70 reduction in inequality has been overestimated.

In conclusion, it appears that the inegalitarian bias in relative consumer price changes in the United Kingdom, which has been a feature now for more than twenty years, has recently accelerated. Before 1970 the price increases in housing were the main element in this bias. However, recently food price increases, especially after the middle of 1972, have taken on a dominant role. Not only does this have the consequence that money measures of inequality are likely to be seriously misleading but that a single consumer price index is no longer adequate for understanding and formulating social policy.

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TABLE V

Effects of Individual Price Increases on Inequality Indices for Various Values of δ [1]

δ.	1970 real *I*.		With 25% price increase in:								
			Food.	Clothing.	Housing.	Fuel and Light.	Drink and Tobacco.	Travel.	Misc. Goods.	Misc. Services.	Durables.
		I	0·1583	0·1473	0·1541	0·1496	0·1515	0·1482	0·1477	0·1521	0·1430
−2·5	0·1495	ΔI	+0·0088	−0·0022	+0·0046	+0·0001	+0·0020	−0·0013	−0·0018	+0·0026	−0·0065
		ΔI*	+50%	−13%	+26%	0%	+11%	−7%	−10%	+15%	−37%
		I	0·1307	0·1218	0·1273	0·1236	0·1252	0·1225	0·1222	0·1257	0·1181
−2·0	0·1236	ΔI	+0·0071	−0·0018	+0·0037	0·0000	+0·0016	−0·0011	−0·0014	+0·0021	−0·0055
		ΔI*	+48%	−12%	+25%	0%	+11%	−7%	−10%	+14%	−37%
		I	0·1016	0·0949	0·0990	0·0962	0·0974	0·0954	0·0951	0·0978	0·0919
−1·5	0·0962	ΔI	+0·0054	−0·0013	+0·0028	0·0000	+0·0012	−0·0008	−0·0011	+0·0016	−0·0043
		ΔI*	+47%	−11%	+24%	0%	+10%	−7%	−10%	+14%	−37%
		I	0·0706	0·0660	0·0688	0·0669	0·0677	0·0663	0·0661	0·0680	0·0638
−1·0	0·0669	ΔI	+0·0037	−0·0009	+0·0019	0·0000	+0·0008	−0·0006	−0·0008	+0·0011	−0·0031
		ΔI*	+47%	−11%	+24%	0%	+10%	−7%	−10%	+14%	−39%
		I	0·0369	0·0345	0·0360	0·0350	0·0354	0·0347	0·0346	0·0356	0·0334
−0·5	0·0350	ΔI	+0·0019	−0·0005	+0·0010	0·0000	+0·0004	−0·0004	−0·0004	+0·0006	−0·0016
		ΔI*	+48%	−12%	+25%	0%	+10%	−7%	−10%	+15%	−40%

[1] These percentages should be raised by a factor of about 1·15 if the differences in the rates of non-response in the FES in 1964 and 1970 are taken into account.

Appendix 1

SOME MEASUREMENT PROBLEMS

A. *Some Difficulties in the Linear Expenditure System*

As was pointed out above, the LES, like all demand systems, is defined only for non-negative demands. This raises two possible problems: of estimation bias and of defining cost-of-living indices. The aggregation consistency of the LES which is due to its linearity breaks down if for some consumers the implied demands are negative. This is illustrated in the figure below:

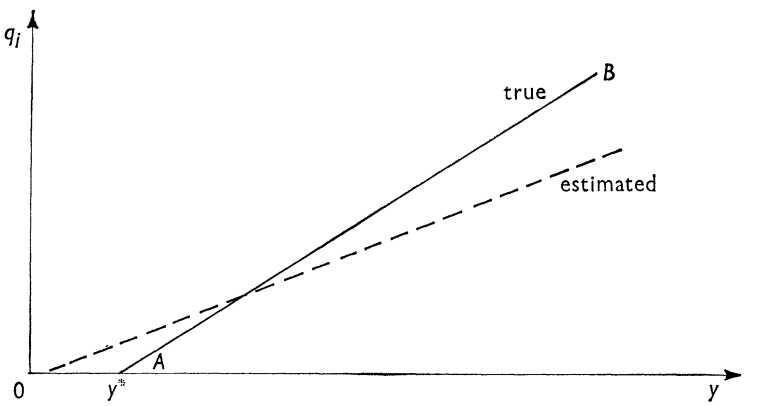

Fig. 1.

AB is the Engel curve for good i defined for $q_i \geqslant 0$, *i.e.*,

$$\alpha_i + \frac{\beta_i}{p_i}(y - \Sigma\, \alpha_j p_j) \geqslant 0$$

i.e.,

$$y \geqslant \Sigma\, \alpha_j p_j - \frac{\alpha_i p_i}{\beta_i} = y^*$$

If some consumers have lower total expenditure than y^*, the model does not apply; *i.e.*, they purchase none of good i. Then the *estimated* demand function will have β_i, which is the slope of AB, underestimated and probably α_i overestimated. The only good for which this is likely to be a problem in practice is durables. From the 1964 equivalent adult expenditure distribution, I have estimated that about 16% of the equivalent adult population had a total expenditure less than that necessary to make durables demand non-negative on the estimated parameters. In 1970 the proportion was about 9%. Since β_i for durables is probably underestimated and α_i overestimated and since durables have had well below average inflation, this means that the cost-of-living index for the well to do has been overestimated. Conversely it is easy to show that for the poor with total expenditure above y^*, the cost-of-living is underestimated. For those with expenditure below y^*, an increase in the price of durables, *ceteris paribus*, would decrease the cost-of-living as defined over 9 goods. This follows from the well-known result

$$\frac{\partial m(p, u)}{\partial p_i} = q_i \quad \text{and} \quad q_i < 0 \quad \text{for } y < y^*$$

This problem has, however, been taken care of by redefining over the remaining 8 goods the cost-of-living and real income indices of those with total expenditure below y^*. This was discussed in Section III above.

It can be argued that durables are somewhat problematical anyway. One can argue, even more strongly than for non-durables where there might for example be habit effects, that there should be a dynamic element in the durables equation. However, there is some evidence that the dynamic specification is reasonably adequate.[1]

There remains the point that from the welfare point of view durables' services and rentals rather than purchases and prices are the relevant arguments in the direct and indirect utility functions respectively. However, if new durables' prices move in parallel with used durables' prices and second-hand markets are well developed and deterioration rates are fairly constant over time, then the stock and the service prices will move in parallel. Given the above point about stock-adjustment, one can argue that the estimated parameters would be similar if data on durables' services and rental prices were available.

There are likely also to be problems with the way household composition effects were incorporated in the LES. In this paper I have assumed that scale economies in household size affect all goods proportionately. As we saw in Section III this had great advantages in simplicity. However, full treatment of the matter will have to be postponed. A pilot investigation for housing led to inconclusive results on the likely direction of bias stemming from this assumption.

The LES is not the only system of demand equations which might be hypothesised. My guess is that similar results would have been obtained with a different functional form which permitted differences in the income responses. Finally, the fact that it was estimated on time series rather than cross-section data may have caused some problems. I hope one day to re-estimate the LES on cross-section data. This is a major undertaking.

B. *The Breadth of the Commodity Groups*

The commodity groups for which the LES was estimated were rather broad. Implicitly it has been assumed so far that *within* each group different income groups either have similar expenditure patterns or that prices for different goods in any given group moved similarly. Fortunately there is some rather strong evidence on this question. The Department of Employment Gazette (see February 1973) publishes price indices for one- and two-pensioner households as well as the General Index of Retail Prices, all excluding housing. Since pensioners have substantially different expenditure patterns from the average—in large measure because of their lower incomes—one would expect within-group differences to emerge in different category price indices than in the General Index. These indices are shown in Table VI. For only three groups do the indices differ by more than about 3% over 1963 to 1971. One of these is an unimportant item,[2]

[1] Ken Wigley has very kindly made available to me some estimates of some stock adjustment models of durables demand equations disaggregated into vehicles, furnishings and floor coverings and other household durables. The hypothesis that there is no stock adjustment behaviour in each separately cannot be rejected on his evidence.

[2] The weight in 1971 in the General Index excluding housing was only 0·05 and 0·019 and 0·007 for one and two pensioner households.

TABLE VI

Price Indices: Annual Averages for Selected Years

(January 16, 1962 = 100)

Year.	All items (excluding housing).	Food.	Alcoholic drink.	Tobacco.	Fuel and light.	Durable household goods.	Clothing and footwear.	Transport and vehicles.	Misc. goods.	Services.	Meals bought and consumed outside the home.
					Index for one-person pensioner households.						
1964	107·0	107·5	108·6	105·8	108·5	100·5	104·7	111·6	106·4	105·0	108·1
1970	140·2	138·2	143·9	136·9	146·8	124·7	120·8	156·9	145·3	148·3	143·6
1971	154·4	153·9	152·0	139·1	161·8	133·3	129·0	189·3	161·5	160·8	160·7
1972	166·2	167·5	158·4	140·1	175·3	138·0	138·2	203·0	172·7	170·6	176·2
					Index for two-person pensioner households.						
1964	107·2	103·1	108·2	105·9	108·3	101·7	105·3	109·1	106·2	103·8	108·1
1970	140·3	139·7	144·7	137·3	147·2	127·7	123·8	151·7	141·4	145·4	143·6
1971	154·2	155·3	154·2	139·5	162·6	137·0	132·3	175·1	157·3	159·3	160·7
1972	165·6	169·7	160·9	140·5	176·1	141·3	141·6	187·1	167·5	160·8	176·2
					General index of retail prices.						
1964	106·2	107·8	107·9	105·8	109·3	102·3	104·9	102·1	105·0	106·9	107·5
1970	138·1	140·1	143·9	136·3	145·7	126·0	123·8	132·1	142·8	153·8	145·5
1971	151·2	155·6	152·7	138·4	160·9	135·4	132·2	147·2	159·1	169·6	165·0
1972	161·2	169·4	159·0	139·5	173·4	140·5	141·8	155·9	168·0	180·5	180·3

Source: Department of Employment Gazette, February 1973, Table 5.

meals outside the house. Of the two remaining, services increased in price rather less for pensioners but the transport and vehicles index increased sharply. On closer examination this proved to be because the public transport price index has increased quite dramatically: Dec. 1972 = 213 compared with the motoring and cycling price index Dec. 1972 = 143 (Jan. 1962 = 100).[1] While it is true that transport and vehicles has a much smaller weight in the pensioner indices than in the General Index, the difference is much less for the working population. Thus, it is quite obvious that *within* the travel category there has been a price change which has injured the poorer members of the working population much more than the better off. There can be little doubt that this completely overshadows the likely fact that within the service group, the services consumed by the less well off have increased in price less (for pensioners over 1964–1972 the increase was 9–10% less than average). Once again then, it is clear that the bias in my measurements *understates* the inegalitarian effect of relative price changes.

Discussion of housing which is omitted from Table VI is postponed to the next section.

C. *Some Problems with the Family Expenditure Survey*

The two areas of concern here are (*a*) the treatment of housing costs and (*b*) the question of systematic variations in the response rate.

(a) *Housing*

The treatment in the FES of housing costs for non-owner occupiers is quite straightforward. They are defined primarily as rent and rates minus any subletting receipts. For owner occupiers, however, weekly housing costs are defined as the weekly equivalent of rateable value (an imputed rental value equivalent) plus actual rates plus repairs, insurance payments minus receipts for subletting. Mortgage payments, outright purchases and major alterations are not included in housing costs or in total expenditure on goods and services. The imputed rental is defined to be some fixed proportion of rateable value scaled by a price index for rented accommodation.[2]

There are a number of serious difficulties with this. No distinction is made between those owning outright and those still paying off their mortgages. This also implies that when mortgage rates increase, the measured distribution *ceteris paribus* between these categories remains unchanged.

In principle, to increase imputed rentals according to an index of rents makes sense: housing services are valued by the return from giving up one unit and renting it out. However, the housing market is notoriously far from being a perfect market. Not only are information costs high but extra-market allocative mechanisms are important—especially for the controlled part of the private rented sector and for council housing. There seems to be evidence to suggest that the private rented sector has shrunk and part of the excess demand for rented accommodation has no doubt been shifted to house purchase. For the controlled part, " key

[1] Source: *Monthly Digest of Statistics*, January 1973, Table 172.

[2] Column 5 in Table VII is a good indicator of rents. It is the housing component of the Index of Retail Prices. In 1972 rents, rates and repairs, etc. had respective weights of 54%, 32% and 14%. An index of rents was published only from 1968 and on a 1962 base was only 1% different in 1972 from this housing price index.

money " and fees for " furniture and fittings " have escalated. It seems likely, therefore, that an index of average rents significantly understates the marginal rental which could be obtained by an owner-occupier giving up a unit of housing services. An index of house prices is thus arguably more relevant for adjusting imputed rents. As Table VII shows, such an index (or indices) show more inflation over 1964–70 than rents.[1]

TABLE VII

Various Price Indices Associated with Housing

	(1) Land.	(2) New houses.	(3) Existing houses.	(4) " Housing."[a]	(5) " Housing."[b]	(6) Council housing.
1963	100·0	100·0	100·0	100·0	100·0	100·0
1964	113·5	107·6	117·2	106·7	105·2	105·6
1965	127·0	117·7	131·5	113·8	111·2	112·8
1966	135·1	126·6	135·7	121·1	118·5	124·0
1967	137·8	134·2	144·2	125·3	124·1	135·2
1968	159·5	141·8	156·8	130·0	130·4	149·6
1969	198·6	151·9	164·3	136·9	135·6	162·4
1970	202·7	160·8	172·5	146·9	145·8	181·6
1971	250·0	181·0	196·2	160·5	159·2	198·4
1972	412·2	232·9	260·1	174·0	175·9	216·8

Identification

(1) price/plot of private sector housing land
(2) average price of new dwellings from building society data
(3) average prices of existing houses mortgaged to Nationwide Building Society, recorded in June of each year
(4) " housing "[a] from Table I above
(5) " housing "[b] component of General Index of Retail Prices
(6) average weekly rents of local authority dwellings in England and Wales, in April of each year

Sources

(1), (2), (6) from *Housing and Construction Statistics*, No. 4, 1972, Department of the Environment, H.M.S.O.
(3) data supplied by Nationwide Building Society
(5) from *Monthly Digest of Statistics*, H.M.S.O.

A quite separate but very important point is that the FES expenditure (and income) data exclude capital gains. It is not obvious why if an imputation of housing services can be allocated to owner occupiers, the same could not be done for capital gains on housing. My inequality measures ignore these very substantial capital gains.[2]

A related dimension of the problem is the following: let " being an owner-occupier " be the relevant good. One of the characteristics of this good is the tax advantage which results. As a result of the dramatic rises in house prices and interest rates, access to this good has now been substantially restricted. The implied redistribution in an inegalitarian direction has been in part one of the older *vs.* the younger. It is completely ignored in my measures.

[1] However, there may be insufficient allowance for quality improvements, such as central heating, in these data.

[2] Except to the extent that such capital gains may increase a household's permanent income and hence may be reflected, perhaps with a lag, in increased expenditures. See Appendix 2 for discussion of some estimates of the effects on inequality of such capital gains.

For what it is worth, council rents seem to have increased more than private rents over 1964–72.[1] It is not clear, however, what the distributional implications of this are. A proper study would have to treat at least owner occupiers, private tenants and council tenants as separate groups. Apart from the very important omission of capital gains which has benefited owner-occupiers and the availability squeeze on rented accommodation which has hurt the less well off and which has not been fully reflected in official rent indices, it is not clear whether in other respects there is a systematic bias in my estimates of the effects of relative price changes with respect to housing.

(b) *Non-response*

Another problem with the FES concerns variations in response rates. From analysis of data on non-response and rateable values for 1964 to 1970 supplied to me by the Department of Employment, it is clear that in general non-response increases systematically with increases in rateable value for most of the range of the latter, except at the very bottom where the reverse relationship holds. Further, there seems to be some evidence that in 1964 the extremes of the distribution were better represented than in 1970. The best estimate of the effect of this on inequality change suggests an overestimate of the decline in inequality of the same order of magnitude as the effect of relative price changes.[2] By reducing the measured inequality reduction for 1964–70, the estimate of the *relative* importance of price changes in the 1964–70 inequality change is increased by 1–2%.

Appendix 2

THE EFFECTS ON INEQUALITY OF CAPITAL GAINS IN HOUSING

The analysis of the effects of consumer price changes has so far been net of any capital gains. Part of the reason for segregating this topic into an appendix is in the rather controversial nature of the topic, and part in the crudity of the assumptions which were necessary to evaluate the impact in the absence of cross-classified expenditure/housing data in the FES.

Following the Hicksian definition, income can be defined as that amount of consumption expenditure which will leave net worth unaffected. Thus income = consumption plus the change in net worth. To convert to real terms, consumption should be deflated by a consumer price index and the capital gains component in net worth deflated by the expected change in the consumer price index. A serious difficulty with respect to capital gains was footnoted in Appendix 1C: this is the problem that if the permanent income hypothesis is valid, consumer expenditures will reflect anticipated capital gains. Then only unanticipated capital gains should properly be added to the expenditure data reported in the FES. If the purpose of the study is to get at the distribution of *permanent income*, these " windfall " gains have to be discounted and distributed as an annual flow. This would naturally substantially reduce the impact on the measure of inequality. However,

[1] In the G.L.C. area, the increase in council rents was even greater. This problem of regional differences in prices has never been properly investigated and could have significant repercussions on aggregate inequality measures.

[2] However, there is more statistical uncertainty here than surrounds the measurement of the latter. Full details are available in an 8-page note available from the author.

if a short run ex-post concept of income is chosen, such capital gains could have a dramatic impact.

The actual proportional annual rates of increase of house prices from Table VI and the cost-of-living from Table III for a well-to-do household were:

	1964–70.	1969–71.	1970–72.
New houses . .	0·067	0·088	0·185
Existing houses . .	0·064	0·089	0·205
Cost-of-living . .	0·045	0·069	0·069

Because of the difficulties mentioned above, I am not prepared to state a precise value for capital gains in housing. Relative to the cost of living, the 1969–71 figure for the change in house prices is 2% higher—the same as over 1964–70. However, over 1970–72 the increase is dramatic: a divergence of at least $11\frac{1}{2}\%$ per annum between the two. The consequences for inequality in those years must have been dramatic.

I have investigated the effect on real inequality in 1970 of hypothetical annual capital gains in housing of 5% and 10%, making some rather crude assumptions:

(*a*) only the top 40% of the adult equivalent distribution in 1970 are owner-occupiers, *i.e.*, those with an annual expenditure of £715 or greater;

(*b*) an adult equivalent with expenditure of £715 lives in a house with value per adult equivalent of £3 × 715 = £2145;

(*c*) the house values of those with higher expenditure increase with expenditure in a way which can be predicted from the housing equation of the LES.

The assumptions overstate the division between owner-occupiers and the rest but probably understate the value of the houses of the rich because inheritance and other factors are ignored in assumption (c).

The implied capital gains in housing at different levels of 1970 annual expenditure per adult equivalent are:

	£	£	£	£
Expenditure levels.	715	1,000	1,500	2,000
Capital gains at 5%	107·3	128·2	164·9	201·7
Capital gains at 10%	214·5	256·4	329·9	403·3

Even a 5% capital gain in housing has dramatic consequences on real inequality indices for 1970 as can be seen from Table VIII. In fact a $2\frac{1}{2}\%$ capital

TABLE VIII

Inequality Indices for Various Values of δ

δ.	1964 Money.	1970 Real.	1970 Real at 5% capital gains.	1970 Real at 10% capital gains.
−2·5	0·1676	0·1495	0·1854	0·2212
−2·0	0·1387	0·1236	0·1544	0·1857
−1·5	0·1080	0·0962	0·1207	0·1460
−1·0	0·0749	0·0669	0·0840	0·1018
−0·5	0·0391	0·0350	0·0438	0·0531

gain in housing would be enough to wipe out the measured inequality reduction over 1964–70. Even in permanent income terms, one can therefore argue that the 1970–72 house price rises had an effect of this magnitude.

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