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TRENDS IN THE SHARES OF TOP WEALTH-HOLDERS IN BRITAIN, 1923–1981‡

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I. INTRODUCTION

This paper re-examines earlier work by Atkinson and Harrison (1978, 1979), which presented a consistent series of estate-based estimates of the distribution of wealth in England and Wales for the years 1923 to 1972,[1] and conducted an econometric analysis of the trends over time in the share of the top 1 percent of wealth-holders.

Such a re-examination is timely for several reasons. First, estate data allowing a considerable extension of our series are now available.[2] Second, these data make it possible to test the forecasting ability of our original econometric specification. Third, there are grounds for modifying this specification, although, as it turns out, the revised formulation performs only marginally better than the original. Finally, the newly estimated equation can be used to examine the likely impact on the distribution of wealth of recent sharp fluctuations in stock and house prices, both of which play a central role in determining the trend of the share of total wealth owned by top wealth-holders.

The paper is organized as follows. Section II presents our updated estimates of the distribution of wealth, covering the years 1923 to 1981. Then, in Section III, we compare the share of the top 1 percent of wealth-holders for the years 1973 to 1981 with the forecasts from our original

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[1] No suitable estate data were collected for years before 1923. There are also years subsequent to 1923 that lack suitable estate data.

[2] The data have retained basically the same form even though Estate Duty was replaced by Capital Transfer Tax in 1975. Dunn and Hoffman argue that this change has not materially affected the coverage of the estate statistics (1983, p. 457).

equation, which was based on data for 1923 to 1972. This assessment of the equation's forecasting performance leads us to propose an alternatively specified equation in Section IV. Finally, Section V concludes the paper by using our equation to reflect on the possible effects on the share of the top 1 percent of the October 1987 stock-market crash and recent increases in house prices.

II. A CONSISTENT SERIES, 1923–1981

The estate multiplier method of estimating the distribution of wealth from estate data uses the estates of those dying in a particular year as a sample of the wealth-holdings of the living population.[3] In a number of respects, a British estate-based sample is not completely representative, and various refinements can be made to overcome this problem. The wealthy tend to live longer, so mortality multipliers (the inverse of the mortality rates) are adjusted according to social class. Not all those who die leave estates that are included in the statistics, requiring an adjustment for what is hereafter referred to as the excluded population. Furthermore, not all the wealth of even the included population appears in the estate data because of the tax treatment of, for example, settled property. These issues are discussed at length in Atkinson and Harrison (1979, pp. 91–94); here we simply describe the main characteristics of the series we have estimated.

- *Period:* 1923–81.
- *Geographical coverage:* England and Wales; Great Britain.[4]
- *Mortality multipliers:* age- and sex-specific multipliers, adjusted by social class differentials obtained from the Registrar General's Decennial Supplement (Registrar General, 1978). No adjustment for social class is made to multipliers applied to smaller estates, defined as those below £10,000 in 1972, £12,500 in 1973 and 1974, £15,000 in 1975 and 1976, £17,500 in 1977, £20,000 in 1978 and 1979, and £25,000 in 1980 and 1981, the amounts being chosen to maintain broadly unchanged the proportions of estates affected.[5,6]
- *Total population:* those economically independent, defined as those

[3] For a fuller discussion of the estate multiplier method, see Atkinson and Harrison (1978, pp. 7–11).

[4] The figures for Great Britain cover the shorter period 1938–81. The availability of estate data for Northern Ireland from 1974 means that estimates for the United Kingdom can be produced for 1974 and later years. As one might expect, the British and UK estimates tend quite closely to reflect movements in the estimates for England and Wales.

[5] Corrections are made to the social class differentials to allow for those classified by the Registrar General as unoccupied and, up to 1972, for errors in occupational statements (the multipliers from 1973, based on the 1971 Decennial Supplement, are not so adjusted since such errors are no longer systematically in a particular direction (Registrar General, 1978, p. 23)).

[6] The social class differential for the estate age category 'age not stated' is calculated as a weighted average of the differentials for those aged 45 and over, using the deaths in 1971 as weights.

above an age threshold reduced linearly from 23 in 1923 to 18 in 1973, and maintained at 18 thereafter.

- *Total wealth:* unadjusted wealth as estimated from the estate data, plus the wealth of that part of the economically independent population not accounted for by the estate data (the excluded population).[7]
- *Definition of wealth:* no adjustments are made for settled property missing from the estate data, for the value of pension rights, or for the over-valuation of life policies.

The new series of estimates is presented in Table 1. The percentage shares are derived using log-linear interpolation, which is a source of potential error when the range of interpolation is broad; this is not, however, a problem with more recent estimates. Despite our attempts to produce a consistent series, there remains a possible lack of comparability between certain sub-periods covered by the series because of changes in the form and coverage of the estate data. Where this problem is most acute — pre- and post-World War II and between 1959 and 1960 — it is indicated in the table by blanks, and the implications are examined in the econometric analysis of Section III.

The cost of aiming for a consistent series is that our estimates are not always the most accurate that can be produced for any individual year. In particular, the figures for recent years could be refined in several respects: for example, no adjustments are made for problems of valuation or for wealth missing from estate returns other than that of the excluded population.[8] The Inland Revenue estimates of the distribution of wealth in Great Britain for some recent years (1966, 1971 and 1974–83) have, however, been adjusted to take account of some of these shortcomings, following the developments described by Dunn and Hoffman (1978). In the Series C estimates, which are most closely comparable to ours, the Inland Revenue revalues life assurance policies (replacing the maturity value included in the estate by an estimate of the equity value) and consumer durables (replacing the realizable value by written down replacement cost), and makes additions for surviving spouse settlements and other trusts.[9] Additionally, the Inland Revenue now uses mortality multipliers classified by marital status. The estimates in Table 1, therefore, besides relating to England and Wales rather than Great Britain, are in other respects not comparable to the Inland Revenue's Series C figures. Despite this, Table 2, which compares our figures with the revised Series C,[10]

[7] In earlier work, the wealth of the excluded population was based on a range of assumptions, but here we present only the central figure, our preferred estimate, referred to as assumption B3 in Atkinson and Harrison (1978, p. 91). The method of calculation of the wealth of the excluded population is described in an appendix, available from the authors.

[8] These issues are discussed in Atkinson and Harrison (1979, pp. 97–98), as is the likely effect on the degree of concentration and on the trend over time if account could be taken of the deficiencies.

[9] Further adjustments are made to yield Series D, which includes occupational pension rights, and Series E, which includes in addition state pension rights.

[10] Table 2 also presents Series C as originally published; differences between the revised and original series are discussed below.

TABLE 1

Shares in Total Wealth, 1923–81

	England and Wales				*Great Britain*			
	Top 1%	*Top 5%*	*Top 10%*	*Top 20%*	*Top 1%*	*Top 5%*	*Top 10%*	*Top 20%*
1923	60.9	82.0	89.1	94.2				
1924	59.9	81.5	88.1	93.8				
1925	61.0	82.1	88.4	93.8				
1926	57.3	79.9	87.4	93.2				
1927	59.8	81.3	88.3	93.8			(see note 1)	
1928	57.0	79.6	87.2	93.1				
1929	55.5	78.9	86.3	92.6				
1930	57.9	79.2	86.6	92.6				
1936	54.2	77.4	85.7	92.0				
1938	55.0	76.9	85.0	91.2	55.0	77.2	85.4	91.6
1950	47.2	74.3			47.2	74.4		
1951	45.8	73.6			45.9	73.8		
1952	43.0	70.2			42.9	70.3		
1953	43.6	71.1			43.5	71.2		
1954	45.3	71.8		(see note 2)	45.3	72.0		(see note 2)
1955	44.5	71.1			43.8	70.8		
1956	44.5	71.3			44.0	71.1		
1957	43.4	68.7			42.9	68.6		
1958	41.4	67.8			40.9	67.7		
1959	41.4	67.6			41.8	67.9		
1960	33.9	59.4	71.5	83.1	34.4	60.0	72.1	83.6
1961	36.5	60.6	71.7	83.3	36.5	60.8	72.1	83.6
1962	31.4	54.8	67.3	80.2	31.9	55.4	67.9	80.7
1963					(see note 3)			
1964	34.5	58.6	71.4	84.3	34.7	59.2	72.0	85.2
1965	33.0	58.1	71.7	85.5	33.3	58.7	72.3	85.8
1966	30.6	55.5	69.2	83.8	31.0	56.1	69.9	84.2
1967	31.4	56.0	70.0	84.5	81.5	56.4	70.5	84.9
1968	33.6	58.3	71.6	35.1	33.6	58.6	72.0	85.4
1969	31.1	56.1	67.7	83.3	31.3	56.6	68.6	84.1
1970	29.7	53.6	68.7	84.5	30.1	54.3	69.4	84.9
1971	28.4	52.3	67.6	84.2	28.8	53.0	68.3	84.8
1972	31.7	56.0	70.4	84.9	32.0	57.2	71.7	85.3
1973	27.3	50.8	66.8	84.9	27.4	51.5	67.5	85.4
1974	22.6	47.8	64.1	83.1	22.9	48.6	65.0	83.6
1975	22.7	45.8	61.9	80.8	23.1	46.5	62.5	81.1
1976	24.4	48.7	65.1	83.7	24.6	49.0	65.4	84.0
1977	22.1	46.5	62.5	81.0	22.1	46.4	62.5	80.9
1978	21.9	45.6	62.4	81.5	22.0	45.9	62.9	81.9
1979	21.5	45.2	61.2	80.3	21.4	45.3	61.4	80.5
1980	19.4	42.4	59.3	79.4	19.6	42.8	59.8	79.9
1981	22.7	45.9	62.6	82.3	22.5	46.0	62.8	82.5

TABLE 2

A Comparison of Estimates of Shares in Total Wealth in Great Britain

	1966	*1971*	*1974*	*1975*	*1976*	*1977*	*1978*	*1979*	*1980*	*1981*
Our estimates										
Top 1%	31	29	23	23	25	22	22	21	20	23
Top 5%	56	53	49	46	49	46	46	45	43	46
Top 10%	70	68	65	62	65	63	63	61	60	63
Revised Series C										
Top 1%	33	31	23	24	24	23	23	22	20	22
Top 5%	56	52	43	44	45	44	44	40	39	42
Top 10%	69	65	57	58	60	58	58	54	52	56
Original Series C										
Top 1%	33	31	23	24	24	23	23	24	23	23
Top5%	56	52	43	44	46	44	44	45	43	45
Top 10%	69	65	57	58	61	58	58	59	58	60

Sources

Our estimates are taken from Table 1, above; the original Series C is taken from issues of *Inland Revenue Statistics* for 1981, 1982, 1984, and 1985; the revised Series C comprises new estimates for 1976 and 1979–81 (from *Inland Revenue Statistics, 1984* and *Inland Revenue Statistics, 1985*) and the estimates from the original Series C for all other years.

shows the same general impression from both sets of estimates: a fall in the shares of top wealth-holders in the first half of the 1970's, followed by a much slower downward trend. This is true of the share of both the top 1 percent and top 5 percent, indicating a persistent feature of the distribution of wealth: the relative constancy of the share of the 4 percent of top wealth-holders immediately below the top 1 percent. In 1923, this group owned 21.1 percent of the personal wealth in England and Wales; in 1981, the equivalent figure was 23.2 percent.

The shares of the top 10 percent and 20 percent, like those of the top 1 percent and 5 percent, both changed little from 1974 onward, but between 1970 and 1974 they behaved noticeably differently. Taking the period 1970–81 as a whole, the share of the top 10 percent in the England and Wales distribution declined roughly in line with the shares of the top 1 percent and 5 percent — by 6.1 percentage points, compared to 7.0 (top 1 percent) and 7.7 (top 5 percent) — but the share of the top 20 percent fell by only 2.2 percentage points. The major beneficiaries of the reduction in the

NOTES to Table 1

1. Estate data for Scotland are not available before 1938.
2. For the years 1950–59, the range of the estate data does not allow estimates of the shares of the top 10 percent and 20 percent.
3. The estate data were not available by country for 1963 so that it was not possible to calculate estimates comparable with those for other years.

share of the top 1 percent were, therefore, those in the range from the sixth to the twentieth percentile, whose share increased by 5.5 percentage points. This group has been regularly improving its position since 1960, benefiting from falls in the shares of the top 5 percent to the point where almost no improvement in the share of the bottom 80 percent is evident. In 1960, the share of the bottom 80 percent was 16.9 percent; by 1981, this figure had risen only to 17.7 percent. This may in part reflect the omission of pension rights from the definition of wealth, but even so this observation tempers any inclination one might have to proclaim a general levelling up of the distribution of wealth on the basis of behaviour of the share of the top 1 percent.

Some readers might wish to take issue with the conclusion we draw from Table 2 that our estimates and the revised Series C describe broadly similar movements in the shares of top wealth-holders. Most notably, the fall between 1971 and 1974 for the top 10 percent is much greater in Series C (8 percentage points) than in Table 1 (3 percentage points), and between 1975 and 1978 there is typically a 5 percentage point difference in the estimated shares of the top 10 percent, widening to 7–8 points from 1979. There is, however, much less divergence between our estimates and the original Series C figures, also shown in Table 2, from which it is clear that the revisions to the Inland Revenue series have appreciably contributed to the differences between Table 1 and Series C.

When revisions have been made, they have typically not been applied to Series C estimates for all years. In 1984, there was a re-appraisal of the estimate of the amount of jointly-held property, and the figures for 1976 and 1979 to 1981 were revised, but not those for other years (*Inland Revenue Statistics 1984*, p. 43). In 1985, there was a less important revision relating to the methods used to adjust the estimates to a balance sheet basis, but revised figures were given only for 1980 to 1982 (*Inland Revenue Statistics 1985*, p. 51). No revisions have been made to estimates for 1977 and 1978, nor to those for 1975 and earlier years. In consequence, although the revisions reflect improvements in the methodology, the resulting estimates for different years are not derived on a consistent basis. For this reason we feel that the series in Table 1, less satisfactory than the Inland Revenue estimates for some individual recent years though it is, nevertheless provides a firmer basis for the examination of the long-term *trend*.

III. FORECASTING THE SHARE OF THE TOP 1 PERCENT, 1973–81

In earlier work, Atkinson and Harrison (1978, 1979) derived a model of the behaviour of the share of the top 1 percent in the total wealth distribution from a simplified version of the process analysed by Meade (1964), incorporating the key factors believed to affect the differential rate of accumulation by the top 1 percent.

The accumulation model shows the wealth, K, of a particular group as governed by the differential equation

$$\frac{\mathrm{d}K}{\mathrm{d}t} = \theta K + \sigma$$

where θ denotes the rate at which wealth tends to reproduce itself, or the 'internal rate of accumulation', and σ denotes saving out of other forms of income. Integrating, we obtain

$$K(t) = K(0)\, e^{\theta t}\left[1 + \frac{\int_0^t \sigma e^{\theta(t-u)} \mathrm{d}u}{K(0)\, e^{\theta t}}\right].$$

For the population as a whole, we assume that σ is dominated by savings out of earnings, and that the second term in the square bracket can be represented by $\alpha_2 PW$, where PW is the ratio of the value of consumer durables and owner-occupied housing (hereafter referred to as popular wealth) to the value of other wealth, and the coefficient α_2 takes account of the proxy nature of PW. For the top 1 percent, the same differential equation is assumed to apply, with θ_1 and σ_1 replacing θ and σ. Integrating produces an equation for $K_1(t)$ that is equivalent to that for $K(t)$; the second term in the square bracket in this case is assumed to be related to the accumulated capital gains on shares (saving out of earnings being assumed to be relatively unimportant for this group), and is therefore represented by $\alpha_3 \pi$, where π is the index of share prices.

We therefore have

$$K(t) = K(0)\, e^{\theta t}[1 + \alpha_2 PW]$$

and

$$K_1(t) = K_1(0)\, e^{\theta_1 t}[1 + \alpha_3 \pi].$$

The equation for the *share* of the top 1 percent is obtained by taking natural logarithms, subtracting $\log K$ from $\log K_1$, and approximating $\log[1+x]$ by x. This yields the linear regression equation:

$$\omega = \alpha_0 + \alpha_1 T + \alpha_2 PW + \alpha_3 \pi + \varepsilon$$

where ω is the logarithm of the share of the top 1 percent and T is a time trend. The model predicts $\alpha_2 < 0$ and $\alpha_3 > 0$; α_1, which equals $\theta_1 - \theta$, the difference between the rates of accumulation, is of indeterminate sign because of opposing considerations.[11] Two dummies were added to the equation to test for the possible breaks in the series mentioned earlier ($D_1 = 0$ for pre-Second World War years, and 1 thereafter; $D_2 = 0$ until 1959, and 1 thereafter), and the modified equation was estimated using the data in Table 1 for England and Wales from 1923 to 1972.

[11] If the top 1 percent receive a higher gross rate of return, α_1 may be positive; if they are taxed more heavily, α_1 may be negative.

We re-estimated this equation[12] and discovered two minor errors in the result as originally reported in Atkinson and Harrison (1979, p. 105).[13] The corrected result is given as equation (3.1) in Table 3. Absolute t-ratios are in parentheses, and the Durbin-Watson statistic, which reveals that the test for first-order serial correlation is inconclusive, is adjusted for missing observations (Savin and White, 1978). The Q-statistic proposed by Ljung and Box (1978) is also reported, as a test for general serial correlation of the residuals. Q is treated as χ^2 with M degrees of freedom, where M is selected according to $M = \min(N/2, 3\sqrt{N})$, in which N is the number of observations. The figure in parentheses beneath the Q-statistic is its critical level, the significance level at which the null hypothesis is just rejected (Lehmann, 1959, p. 62). In the case of equation (3.1), therefore, the null hypothesis cannot be rejected at any conventional significance level.

The predicted value of the share of the top 1 percent in 1972 using this equation is 31.1 percent. From Table 1, it is clear that this is quite close to the actual figure of 31.7 percent, but the table also indicates that there have been substantial changes in the share since that date, taking it well outside the range for the period of estimation. The figure for 1981, for example, is fully one-fifth lower than the lowest value recorded in the period over which the equation was estimated. It is for this reason that we begin by examining the forecasting performance of the equation, rather than re-estimating with the additional years' data points. Note that there was no possibility to revise the equation's original specification in the light of its forecasting performance over this period: when the equation was estimated, some ten years ago, the additional data were not available.

Consider first the various elements that contribute to the forecast change between 1972 and 1981:

- The time trend reduces the share to 0.923 times its previous value, to 28.7 percent, so that the past trend cannot explain much of the observed fall.
- The expansion of popular wealth, relative to other forms of personal wealth, as a result of the spread of owner-occupation and the rise in house prices, implies a very large reduction of nearly 12 percentage points in the share of the top 1 percent.[14]
- This predicted reduction is partly offset by the rise in share prices

[12] For the empirical work in this paper, we used the econometric software package PC-GIVE (Hendry, 1986).

[13] The Durbin-Watson statistic did not make the correct adjustment for gaps in the series, and there was a misprint in one of the t-ratios.

[14] The 1972–81 figures for popular wealth (defined as the value of housing plus consumer durables) and total wealth are obtained from the official balance sheet estimates (further details of which are provided in the appendix available from the author), and linked to the previous series via overlapping figures for 1972. In each case we take the average of end-year figures to give a figure for the year in question; for example, that for 1973 is based on balance sheet figures for 31 December 1972 and for 31 December 1973.

TABLE 3

Regression Results, Logarithmic Specification

Eqn. No.	*Country*	*Years*	*Constant*	*Time*	D_{-1}	D_{-2}	π	*PW*	R^2	*DW*	*Q*
3.1	EW	1923–72	4.103	−0.00891	−0.0167	−0.236	0.00141	−0.152	0.985	2.446	7.212
				(4.185)	(0.325)	(8.364)	(2.619)	(2.833)			(0.951)
3.2	EW	1923–81	4.111	−0.00892	−0.0194	−0.217	0.00089	−0.114	0.989	2.391	20.203
				(4.186)	(0.356)	(7.997)	(3.766)	(9.964)			(0.322)

NOTES:

1. Figures in parentheses beneath coefficients are absolute t-ratios.
2. Q refers to the Ljung-Box Q-statistic. The figure in parentheses beneath each Q is the critical level.

towards the end of the period, which yields an increase of 5 percentage points in the share of the top 1 percent.[15]

Taken together, these result in a predicted share for the top 1 percent in 1981 of 20.7 percent. This is rather lower than the actual figure in Table 1 of 22.7 percent, but suggests that the earlier model does allow for the possibility of substantial changes in the share of the top 1 percent outside the range of variation during the estimation period; it therefore seems to be a reasonable starting point for further analysis.

A more formal approach to the forecasting performance is provided by considering the predictions for the period 1973–81 as a whole and applying a χ^2 test. The forecast error in 1981, expressed in terms of logarithms, is 0.091, which is 2.65 times the estimated standard error. The values for 1975 and 1976 are larger still, and the sum of squared deviations, normalized by the standard error, is 53.9. This statistic, which would be distributed as χ^2_9 in large samples if the parameters remained constant (Hendry, 1980), leads us to reject the hypothesis of parameter stability, or one-period ahead forecast accuracy.

This asymptotic test may, however, be too severe in the present context. Pesaran, Smith and Yeo note that, for small samples, it 'will tend to over-reject when the null hypothesis is true' (1985, p. 290), and Kiviet's Monte Carlo evidence confirms this (1986, p. 249). Kiviet's recommendation is to use instead the version of the Chow test designed for the case when the forecasting period has too few degrees of freedom to allow a separate regression to be estimated. The test yields a statistic of 2.19, uncomfortably close to the 5 percent critical value of 2.26.

This test is actually one of predictive failure: rejection occurs if either the coefficients change or the errors are heteroscedastic (Pesaran, Smith and Yeo, 1985, p. 288). To investigate the possibility of heteroscedasticity further, we re-estimated the equation for the full data period 1923–81 (equation (3.2)); this produced an increase in the variance of the residuals towards the end of the data period. A separate test for heteroscedasticity (White, 1980) revealed, however, that the null hypothesis of homoscedastic errors could not be rejected, although this may say more about the sample size than anything else.

Judged overall, then, these tests seem to indicate only qualified support for the specification of the estimating equation: it is strongly rejected by the asymptotic test, and almost rejected by the Chow test. We therefore consider next an alternative formulation.

IV. A REVISED SPECIFICATION

The modification we adopted was a simple transformation to the dependent variable. That originally employed, the logarithm of the share of the top 1

[15] The share price index is the Financial Times index for 500 industrial shares published regularly in *Financial Statistics*.

percent, is limited in its range. An alternative log-logistic formulation, which allows variation over an unlimited range, is to take the logarithm of the share of the top 1 percent relative to that of the bottom 99 percent. This formulation remains consistent with the theoretical model from which the estimating equation is derived;[16] additionally the unbounded nature of the dependent variable makes the use of a linear trend more defensible than it is in the original specification.

Equation (4.1) of Table 4 shows the result of adopting the log-logistic specification for the data for England and Wales from 1923 to 1972. The forecasting performance of this equation, judged by the same criteria as those used previously, improves appreciably on that of equation (3.1). Hendry's asymptotic test now yields a statistic of 17.6, slighly above the 5 percent critical value of 16.9 and well below the 1 percent value of 21.7. Bearing in mind the tendency of this test to over-reject in small samples, this figure suggests that predictive failure is not a problem with the revised formulation, and the Chow test statistic of 1.30 firmly supports this conclusion.

These results could be interpreted as conclusive evidence in favour of the log-logistic formulation, but there are grounds for arguing that the tests exaggerate its superiority over the logarithmic specification. Figure 1, which shows the estimates from Table 1 of the share of the top 1 percent, for the years 1973 to 1981, together with the predictions from equations (3.1) and (4.1), suggests that the estimated models track movements in the share of the top 1 percent about as well as one another. How well can be gauged from Table 5. Estimating each equation for 1923–81, but with dummy variables for each of the years 1973–81, is equivalent to estimating the equation for 1923–72, except that it additionally yields the period-by-period prediction errors for 1973–81 (the coefficients on the dummies), given in Table 5 together with the coefficients' absolute t-ratios, which 'test whether each prediction error…differs significantly from zero' (Pesaran, Smith and Yeo, 1985, p. 287). The coefficients reveal that both models have problems with 1976 and, to a lesser extent, 1975; the t-ratios demonstrate, however, that none of the forecast errors for either model is significantly different from zero.[17]

The results of re-estimating with the transformed dependent variable for the full period are given as equation (4.2). Comparing this with the equation for 1923–72, we can see that the additional years' data appear to have had some impact on the coefficients, although it is only those for π and PW that are noticeably affected, the former falling by about a third and the latter by about a quarter; equations (3.1) and (3.2) reveal that an equivalent effect is observed in the case of the logarithmic formulation.

[16] Since the share of the top 1 percent is the ratio of two values of K, the log-logistic formulation simply requires reinterpreting the denominator as applying to the bottom 99 percent, rather than the whole population.

[17] The test that all the dummies in a model are jointly zero is the Chow test already reported (Pesaran, Smith and Yeo, 1985, p. 287).

TABLE 4

Regression Results, Log-Logistic Specification

Eqn. No.	4.1	4.2	4.3	4.4	4.5	4.6	4.7
Country	EW	EW	EW	EW	GB	GB	GB
Years	1923–72	1923–81	1950–81	1923–81	1950–81	1952–81	1952–81
Constant	0.417	0.425	0.394	0.435	0.369	0.223	0.233
Time	−0.0192	−0.0183	−0.0194	−0.0206	−0.0187	−0.0161	−0.01534
	(5.308)	(5.577)	(3.500)	(19.427)	(3.546)	(2.634)	(2.312)
D_1	−0.0439	−0.0616					
	(0.503)	(0.731)					
D_2	−0.354	−0.323	−0.317	−0.308	−0.302	−0.399	−0.381
	(7.401)	(7.720)	(5.742)	(8.504)	(5.833)	(5.780)	(6.301)
π	0.00211	0.00135	0.00141	0.00141	0.00130	0.00173	0.00204
	(2.295)	(3.731)	(3.608)	(4.049)	(3.473)	(2.601)	(4.272)
π_{-1}						0.000738	
						(0.804)	
π_{-2}						−0.00162	−0.00113
						(1.912)	(2.035)

PW	−0.167	−0.138	−0.136	−0.131	−0.134	−0.156	−0.140
	(1.831)	(7.796)	(6.010)	(8.613)	(6.297)	(2.991)	(3.034)
PW_{-1}						0.145	0.117
						(1.641)	(1.484)
PW_{-2}						−0.201	−0.179
						(2.488)	(2.428)
DV_{-1}						−0.286	−0.209
						(1.66)	(1.503)
DV_{-2}						0.0247	
						(0.166)	
R^2	0.986	0.990	0.980	0.990	0.980	0.984	0.997
DW	2.415	2.390	2.322	2.356	2.139		
Q	5.475	14.247	12.859	13.581	18.286		
	(0.977)	(0.712)	(0.613)	(0.756)	(0.248)		

NOTE: (see also Notes to Table 3)

1. DV_{-1} and DV_{-2} refer to the lagged dependent variables.

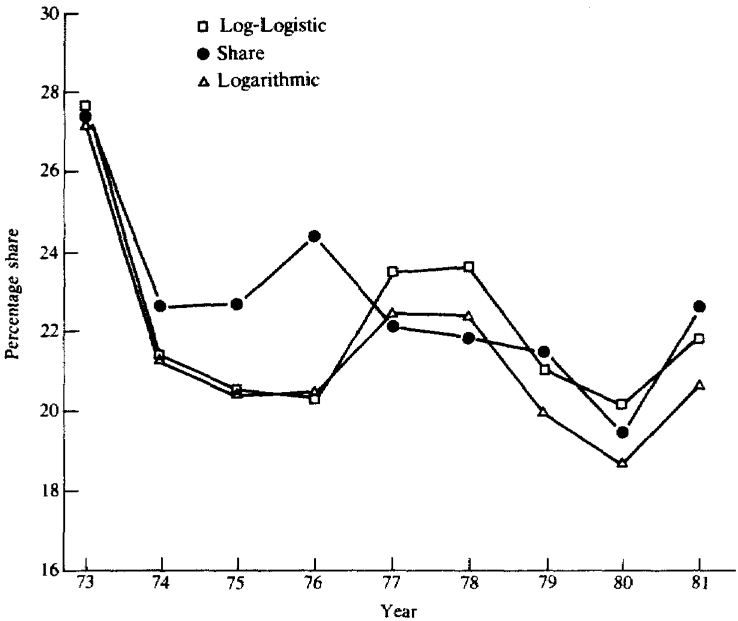

Fig. 1. Actual and predicted shares

There remains the question of whether the dynamic specification of the estimating equation is satisfactory, and this is the final issue we consider. There are good reasons for testing for lagged effects of explanatory variables. A rise in share prices may induce portfolio changes which magnify, or attenuate, the first-round impact. There are also delays that arise in recording the estate data used to generate the estimates. These data cover estates on which duty is first paid in the year ending in the following March. This allows for a lag of 3 months between death and first payment of duty, but to the extent that the delay is shorter, or more likely longer, we are observing wealth covering a different period. The position is further complicated by the treatment of corrections to the original returns, which are entered in the year in which they are made.

If we confine attention to the log-logistic formulation, there is little in equation (4.2) to indicate which lags are likely to be important. The Durbin-Watson statistic delivers an inconclusive verdict on the null hypothesis of the absence of first-order serial correlation. The Q-statistic, for the test of general serial correlation, does not reject the null hypothesis, but the validity of Q-statistics is disputed by Breusch and Pagan who refer to them as '"portmanteau statistics" of dubious power' (1980, p. 244) and advise testing specifically those orders of autoregression that are likely to arise with the data. In the present case, however, we have no strong prior expectations.

TABLE 5
Period-by-Period Prediction Errors and Absolute t-Ratios, 1973–81

	Model			
	Logarithmic		*Log-Logistic*	
Year	*Error*	*t-Ratio*	*Error*	*t-Ratio*
1973	0.0102	0.1850	−0.0104	0.1111
1974	0.0657	0.5343	0.0576	0.2757
1975	0.1104	0.8204	0.1019	0.4459
1976	0.1748	1.3415	0.1824	0.8237
1977	−0.0138	0.1361	−0.0605	0.3507
1978	−0.0232	0.2268	−0.0781	0.4483
1979	0.0765	0.5419	0.0203	0.0848
1980	0.0385	0.2386	−0.0381	0.1391
1981	0.0910	0.6949	0.0351	0.1577

We began therefore with a general model, incorporating two lags on the dependent variable and each of the independent variables. To do this, we had to confine our analysis to a series without gaps. Rather than use a sample beginning as late as 1964, after the last gap, we interpolated a figure for Great Britain for 1963 using the Inland Revenue estimates for that year[18] and thereby generated a continuous series for Great Britain for the 32 years from 1950 to 1981.

Before employing this series as the dependent variable, we first examined the effect of confining attention to the postwar years by estimating an equation with data for England and Wales for the subsample 1950–81 (with one gap in 1963). The result is shown as equation (4.3). If this is compared with the same equation estimated for 1923–81 — equation (4.4)[19] — the F test for structural stability does not reject the null hypothesis. Equation (4.5) is identical to the formulation for equation (4.3) except that it relates to Great Britain rather than England and Wales, with the one gap in 1963 accounted for in the manner described above. The results are very similar, as we would expect given that the estimates for England and Wales and Great Britain in Table 1 differ so little, and support our view that it is reasonable to examine the dynamic structure of our preferred specification with reference to an equation explaining movements in the share of the top 1 percent in Great Britain over the years 1950–81.

[18] See Inland Revenue, *Hundred and Seventh Report*, Cmnd. 2572, 1965, Table 145. The estimates of the size and wealth of the excluded population were taken from Atkinson and Harrison (1978, Appendix VI). The resulting share of the top 1 percent in Great Britain for 1963 is 32.64 percent.

[19] This is the same as equation (4.2) except for the exclusion of D_1, the coefficient on which is in any case insignificant.

The inclusion of lagged dependent and independent variables, reported as equation (4.6), did not lead to significantly improved results. The coefficient of PW_{-2} has a t-statistic in excess of 2, but an F test of the restrictions in equation (4.5) against the unrestricted equation (4.6) yields a statistic of 1.33, compared with a 5 percent critical level of 2.60, so that we cannot reject the hypothesis that the extra coefficients are jointly zero. The estimated equation does, though, suggest a more parsimonious representation, in which the second lag on the dependent variable and the first lag on π are omitted. This results in equation (4.7). A test of the restricted equation (4.5) against equation (4.7) yields an $F(4, 19)$ statistic of 1.96, well below the 5 percent critical level of 2.84. It is still not possible, therefore, to reject the null hypothesis that the extra coefficients, relative to equation (4.5), are jointly zero.

This failure to reject may be a reflection of the smallness of the sample, so that it seems appropriate to consider the error specification nevertheless. Breusch and Pagan (1980, p. 245) argue that, with the inclusion of lagged dependent variables, use of the Q-statistic to test for serial correlation is inappropriate, and this is corroborated by the Monte Carlo study of Kiviet (1986), who suggests use of the Lagrange multiplier type F-test. Testing for first and second order serial correlation, this yields a statistic of 3.37, which is below the 5 percent critical value for $F(2, 19)$. The error specification for this equation appears, therefore, to be satisfactory.

V. CONCLUSIONS AND IMPLICATIONS

In this paper, we present new estimates of the distribution of wealth for the period 1973 to 1981. These estimates are less refined than most of those published by the Inland Revenue, which have been significantly improved over the course of the past ten years. As a consequence of these improvements, however, the official figures do not provide a consistent series over time, whereas the series we have constructed provides estimates on a consistent basis from 1923 to 1981.

This series allows us to examine in more detail the factors lying behind the fall in the share of the top 1 percent of wealth-holders over the past 60 years. We use two formulations for our estimating equation. The first, in which the dependent variable is the logarithm of the share of the top 1 percent of wealth-holders, is equivalent to that originally estimated in Atkinson and Harrison (1978, 1979) for the period 1923–72. The second introduces a log-logistic transformation of the share of the top 1 percent. Both appear to track the marked changes between 1973 and 1981 quite well, but the latter proves superior in forecasting these changes when evaluated by standard tests of forecast accuracy, and is additionally more satisfactory for other reasons.

Both sets of results indicate that a model that explains changes in the share of the top 1 percent in terms of variations in share prices and the ratio of popular wealth to other wealth provides a good fit to the series. According to this explanation, the sharp fall in the early 1970's was due not to any acceleration of the downward trend but to the movement in these variables; the much slower downward movement after that date reflects the recovery of share prices.

In the light of the importance of movements in share prices and popular wealth for the share of the top 1 percent, we conclude by using equation (4.2) to predict the impact of major changes in these variables since 1981. We consider first share prices. Undoubtedly, the most important event in the stock market was the dramatic fall in share prices experienced in October 1987. We assumed a fall in the index of share prices of a quarter, relative to its 1981 value; this yields a reduction in the share of the top 1 percent of approximately two percentage points, a little under one-tenth of its 1981 level.

Popular wealth is defined as the value of consumer durables and owner-occupied housing, and the major influence on this variable in recent years has been the strong upward surge in house prices. To illustrate the effect this has had on the distribution of wealth, we considered a rise in the value of housing of 25 percent (the order of magnitude of annual increase observed in recent years); the predicted effect on the share of the top 1 percent is a reduction of some 2.5 percentage points. Together with the effect of the fall in the stock market, then, we predict an appreciable erosion of the share of the top 1 percent, but one that will have been offset to a degree by the subsequent revival of share prices.

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