CONSUMPTION INEQUALITY AND INCOME UNCERTAINTY*

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This paper places the debate over using consumption or income in studies of inequality growth in a formal intertemporal setting. It highlights the importance of permanent and transitory income uncertainty in the evaluation of growth in consumption inequality. We derive conditions under which the growth of variances and covariances of income and consumption can be used to separately identify the growth in the variance of permanent and transitory income shocks. Household data from Britain for the period 1968–1992 are used to show a strong growth in transitory inequality toward the end of this period, while younger cohorts are shown to face significantly higher levels of permanent inequality.

I. INTRODUCTION

The use of current income in studies of inequality is open to the obvious criticism that current income may not reflect the longer run level of resources available to a household or an individual. Temporarily high or low incomes may exaggerate the true position of the household when borrowing or saving is allowed to smooth the stream of consumption. Moreover, aggregate measures of inequality (or poverty) based on snapshots of income may fail to pick up changes over time associated with the duration rather than the depth of low-income spells. The importance of distinguishing long-run income inequality from inequality associated with transitory movements in income has been emphasized in a number of recent studies that have moved away from simply documenting the change in the cross-section distribution of income. For example, Gottschalk and Moffitt [1994], Moffitt and Gottschalk [1995], Buchinsky and Hunt [1996], and Gittleman and Joyce [1996] all use the time series of individual incomes

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to focus on this important distinction. Moffitt and Gottschalk [1995] study the autocovariance structure of United States male earnings in the Panel Study of Income Dynamics over the 1970s and 1980s. They find a strong increase in the variance of the permanent component in income, mirroring the Buchinsky and Hunt [1996] and Gittleman and Joyce [1996] studies. However, they also show that the increase in income inequality over the later part of this period can be increasingly attributed to a rise in the variance of the transitory component of income, the transitory component showing no significant rise until the early 1980s.

The aim of this paper is to see how consumption can help in this evaluation. The recognition that consumption expenditure may better reflect expected lifetime resources has led to the increasing use of consumption[1] in the measurement of household welfare. Earlier work, such as Cutler and Katz [1991, 1992] and Slesnick [1993], has used repeated cross-section data on the distribution of consumption to examine changes in the distribution of permanent income and household welfare. Attanasio and Davis [1996] exploit the strong systematic movements in wage inequality over the 1980s in the United States to evaluate the complete insurance hypothesis. They find that the distribution of consumption tends to follow the low frequency movements in real wages but that consumption is well insulated from transitory movements. This provides strong evidence against complete insurance while giving tacit support to consumption smoothing.

This study seeks to advance the literature by formalizing some of the limitations of the use of consumption data alone in assessing changes in the distribution of economic welfare and permanent income. We further show how information on changes in the cross-section joint distribution of consumption and income can illuminate the nature of changes in inequality.

As consumption inequality tends to highlight the importance of permanent inequality, the arguments for consumption-based measures of inequality are powerful. But, how reliable is consumption as a measure of welfare? Consumption expenditure does typically differ from income, and these differences surely reflect differences in expected resources and needs. However, although comparisons within date of birth cohorts are likely to be reliable measures of inequality in living standards, this is not so easily

1. In this paper we tend to use the terms “consumption” and “expenditure” interchangeably while recognizing that in practical applications the distinction between the two is of considerable importance.

proved for comparisons across different cohorts. The different intertemporal substitutions open to cohorts at similar points in their life cycle and the different ages at which they are observed make such comparisons less compelling. We therefore argue that there are strong welfare grounds for analysis within cohorts. Moreover, the evolution of distribution within the whole population is influenced by changes in age structure that obscure the role of permanent and transitory income uncertainty.

In this paper we examine the distinction between permanent and transitory income uncertainty in the evaluation of growth in consumption inequality within cohorts. We derive conditions under which the growth of variances and covariances of income and consumption can be used to separately identify the growth in the variance of permanent and transitory income shocks. We develop a difference-in-differences estimator for estimating the growth of the variance of the transitory component on this basis. This is based on a contrast between the growth in the variance of income and consumption. In addition, we show that, where consumption and income are available in the same survey, the covariance between the two provides overidentifying information from which we can verify these results. These results are extended to income processes that include common shocks and cross-sectional correlation between shocks and past incomes. They are also generalized to consumers with preferences that permit precautionary saving. Household data from Britain for the period 1968–1992 are used to show a strong growth in transitory inequality toward the end of this period, while younger cohorts are shown to face significantly higher levels of permanent inequality in comparison to older cohorts at a similar age.

The paper is organized as follows: Section II of the paper is concerned with exploring precisely when consumption does provide a suitable measure of welfare. The theory points to the importance of within-cohort comparisons of consumption and income inequality. Our analysis covers the case of prudent consumers by addressing the relationship between precautionary saving and the welfare cost of risk. Section III considers a variety of income processes and preferences and shows how growth in the variances and covariances of income and consumption can be used to identify growth in the variances of transitory and permanent shocks. These results are then used in Section IV for an empirical evaluation of the differences in growth rates between income and consumption using British Family Expenditure Survey data over the 1970s and 1980s. We provide an analysis by cohort and compare this with the overall picture of changing inequality. Section V concludes.

II. Does Consumption Inequality Measure Welfare Inequality?

A. The Welfare Comparison Case for within-Cohort Comparisons

If one observes an individual's consumption, knowing the individual's age and the interest rates that link the periods of their life, it will always be possible to invert the Hicksian demand function to recover utility given real discount rates and age, assuming consumption in all periods to be a normal good. Given the assumption of common interest rates, comparisons of consumption within cohort at the same point in time do therefore suffice for welfare comparisons. However, if comparisons are between individuals who are differently aged or born in different years, as they will be if comparisons are across cohorts at a given date or across time within a cohort, then comparisons are more problematic.

Suppose that individual i, reaching adulthood in year b_i has lifetime income Y_i. The real interest rate in year s is r_s and is assumed to be the same for all individuals. The individual seeks to maximize an increasing and quasi-concave lifetime welfare function $U_i = U(\mathbf{c}_i)$, where $\mathbf{c}_i \equiv (c_{i0}, c_{i1}, \ldots, c_{iT})$ and c_{it} is consumption at age t, subject to their lifetime budget constraint. Hicksian demands are $c_{it} = c_t(U_i, \mathbf{p}_i)$, where $\mathbf{p}_i \equiv (p_{i0}, p_{i1}, \ldots, p_{iT})$ and $p_{it} \equiv \Pi_{s=0}^{t}(1 + r_{s+b_i})^{-1}$. We assume interpersonal ordinal full comparability of welfare so that welfare comparisons are preserved only by common increasing transformations of utilities.[2]

The following proposition shows the conditions under which consumption comparisons suffice for welfare comparisons.

Proposition 1. (i) *Comparisons within cohorts at same age:* $c_{it} \geq c_{jt}$ implies that $U_i \geq U_j$ whenever individuals i and j share the same year of birth if and only if consumption in all periods is a normal good. (ii) *Comparisons across cohorts at same age:* $c_{it} \geq c_{jt}$ implies that $U_i \geq U_j$ for all i and j whether or not individuals i and j share the same year of birth if and only if $c_t(U_i, \mathbf{p}_i) = f_t(U_i)$, where $f_t(.)$ is an increasing function for all t. This is so if and only if $U(\mathbf{c}_i) = \min_t u_t(c_{it})$ where $u_t(.)$ is an

2. See, for instance, Sen [1977].

increasing function for all t. (iii) *Comparisons across ages:* $c_{it} \geq c_{js}$ implies that $U_i \geq U_j$ for all s and t whenever individuals i and j share the same year of birth if and only if $c_t(U_i,\mathbf{p}_i) = f(U_i)$, where $f(.)$ is an increasing function. This is so if and only if $U(\mathbf{c}_i) = \min_t u(c_{it})$, where $u(.)$ is an increasing function. The same conditions apply to ensure that $c_{it} \geq c_{js}$ implies $U_i \geq U_j$ for all s and t whether or not individuals i and j share the same year of birth.

Proof. See Appendix 1.

In cases of cross-cohort or cross-age comparisons, all requisite information on welfare is available from consumption only if agents choose to equalize utilities across all periods of the life cycle—an extreme case of antipathy to intertemporal substitution. This is obviously an unrealistic degree of smoothing, and as a result there is almost certain to be some weakness in the undiscriminating use of consumption as an indicator of lifetime standard of living.

Additive separability in the direct representation of lifetime utility $U_i = \Sigma_t u_t(c_{it})$ would be a more appealing assumption and one that is commonly adopted. With such preferences the first-order conditions for optimization imply that agents aim for a marginal utility of within-period expenditure equal at each age to the marginal utility of discounted lifetime income:

$$u'_0(c_{i0}) = u'_t(c_{it})/p_{it}. \tag{1}$$

This constancy of discounted marginal utility is the familiar Euler condition for consumption over the life cycle (see Hall [1978], Attanasio and Weber [1989], and Browning, Deaton, and Irish [1985], for example). Even though direct utility is additive across periods, it is known that intertemporal substitution invalidates the use of the sum of compensating variations as a measure of lifetime compensating variation (see Blackorby, Donaldson, and Moloney [1984] and Keen [1990]). Nevertheless, we might wish to categorize the circumstances under which consumption is likely to be reliable as a welfare measure.

Equation (1) points to a number of reasons why consumption could give a poor indication of welfare. If within-period utility functions $u_t(.)$ vary much over the life cycle and welfare comparisons are made across age, then comparisons could be undermined. In particular, subjective discounting will lead to consumption being pushed toward the earlier years of life. On the other hand, if real interest rates are high and welfare comparisons are made across age, then incentives to push consumption toward the later period of life will lead the old to appear to be better off than they actually are in comparison to the young. Finally, if the interest rate is variable over time, then differing incentives to substitute intertemporally could undermine welfare comparisons made across cohorts. The magnitude of this last problem can be shown to depend on the magnitude of intertemporal substitution elasticities.

B. The Welfare Cost of Income Risk

Risk-averse households with more uncertain incomes than others need to be considered worse off. This might show up to an extent in lower expenditure by such households, if their response is precautionary saving (see Kimball [1990] for discussions of precautionary saving). This section explores the precise relationship between precautionary saving and the welfare cost of risk.

Suppose that future income is uncertain and intertemporal utility takes the additive form. It will be useful to define $\tilde{Y}_i$ as that certain present discounted value of income which would allow the individual to achieve the same expected utility. The consumption stream $\tilde{\mathbf{c}}_i = \tilde{\mathbf{c}}(EU_i)$ that would be chosen given $\tilde{Y}_i$ satisfies

$$\sum_t u_t(\tilde{c}_{it}) \equiv E\left(\sum_t u_t(c_{it})\right) = EU_i.$$

Focusing on within-cohort comparisons of $\tilde{\mathbf{c}}(EU_i)$, actual consumption differences will be indicative of welfare differences when households face differing degrees of uncertainty only if $\mathbf{c}_i = \tilde{\mathbf{c}}(EU_i)$ for all households. We show below that the unique case in which this holds exactly is that of constant absolute risk aversion (CARA).[3]

PROPOSITION 2. *Comparisons across individuals facing different income risk:* $c_{it} \geq c_{jt}$ implies that $EU_i \geq EU_j$ whenever individuals i and j share the same year of birth if and only if $\mathbf{c}_i = \tilde{\mathbf{c}}\ (EU_i)$ whatever the distribution of future income. This is so if and only if

$$u_t(c_{it}) = -\alpha_t \exp(-\beta_t c_{it}) \qquad \alpha_t, \beta_t > 0,\ t > 0. \tag{2}$$

3. Strictly speaking, this condition applies to subsequent periods' utility only since period 1 is not affected by uncertainty, but it would seem sensible to maintain the same structure on within-period preferences in all periods.

Proof. See Appendix 1.

The sufficiency part of this is a special case of a more general result of Drèze and Modigliani [1972, p. 324] which also establishes that decreasing absolute risk aversion (DARA) implies that $c_{i0} < \tilde{c}_{i0}$; i. e., that there is excess precautionary saving if higher incomes decrease risk aversion. Their result is not interpreted in terms of welfare considerations but rather in terms of the correspondence between CARA and the absence of substitution effects in response to uncertainty. This links the preference restrictions considered here nicely with those of the previous section. The validity of consumption as a welfare measure relies on the absence of substitution effects in response to intertemporal prices and to uncertainty.

III. Permanent Inequality and Transitory Uncertainty

Uncertainty concerned with unexpected transitory shifts in income is very different from the inequality associated with permanent shifts in the position of individuals in the income distribution. Growth in cross-section measures of income inequality (or poverty) cannot alone distinguish between these two phenomena. Here we show that taking income inequality together with consumption inequality and the life-cycle model, we are able to separate the growth in permanent inequality from the growth in transitory uncertainty.

A. A Stochastic Process for Income

We start by considering a permanent-transitory decomposition for income. Income for individual i in cohort k in period t is written as

$$y_{it} = y^p_{it} + u_{it} \qquad \text{for } i \in k, \tag{3}$$

where y^p_{it} represents the permanent component of income and u_{it} the transitory shock in period t. The permanent component is assumed to follow a random walk:

$$y^p_{it} = y^p_{i,t-1} + v_{it}, \tag{4}$$

where v_{it} is a permanent shock assumed orthogonal to u_{it}. We also assume that the variances of the shocks are the same in any period for all individuals in any cohort but that these variances are not constant over time. The cross-sectional covariances of the shocks with previous periods' incomes are assumed to be zero. In this discussion we assume that shocks are independently distributed across individuals. Subsection II.D below considers a relaxation of this assumption and also considers models for income in which the cross-section distribution of shocks is correlated with the distribution of past income.

The process for income can be written as

$$y_{it} = y_{i,t-1} + u_{it} - u_{i,t-1} + v_{it}. \tag{5}$$

This covers the general MA(1) model in which

$$y_{it} = y_{i,t-1} + \epsilon_{it} - \theta_t \epsilon_{i,t-1}, \tag{6}$$

though notice that in this representation the MA coefficient θ_t is time varying. The evolution of θ_t can be directly related to the evolution of the variances of the transitory and permanent innovations to income.[4] Defining $\mathrm{var}_{kt}(u)$ to be the cross-section variance of transitory shocks for cohort k in period t and $\mathrm{var}_{kt}(v)$ to be the corresponding variance of permanent shocks, the growth in the cross-section variance of income for cohort k can be seen from (5) to take the form,[5]

$$\Delta\,\mathrm{var}_{kt}(y) = \Delta\,\mathrm{var}_{kt}(u) + \mathrm{var}_{kt}(v). \tag{7}$$

Both permanent inequality ($\mathrm{var}_{kt}(v)$) and growth in uncertainty ($\Delta\,\mathrm{var}_{kt}(u)$) result in growth of income inequality. The cross-section distribution of income cannot, on its own, distinguish these.

B. Identifying the Growth in Transitory and Permanent Variances

Taking income inequality together with consumption inequality and the life-cycle model, we are able to separate the growth in permanent inequality from the growth in transitory uncertainty. Assuming quadratic preferences, with the discount rate equal to the real interest rate, we obtain the familiar martingale property for consumption:[6]

$$\rho_t \Delta c_{it} = \eta_{it}, \tag{8}$$

where ρ_t is an annuitization factor[7] and η_{it} is the consumption

4. See Blundell and Preston [1997].
5. Note that $\mathrm{cov}_{kt-1}(yu) = \mathrm{var}_{kt-1}(u)$ where $\mathrm{cov}_{kt-1}(yu)$ is the cross-section variance of income with transitory shocks for cohort k in period $t-1$.
6. See Hall [1978].
7. $\rho_t = 1 - (1+r)^{-(T-t+1)}$.

innovation.[8] Relating η_{it} to the transitory and permanent innovations to income (5), we have

$$\eta_{it} = \rho_t v_{it} + r u_{it}/(1 + r). \tag{9}$$

That is, the consumption innovation is simply the sum of the annuity value of the transitory shock and permanent shock.

The derivation of (9) requires that the consumer can separately identify transitory u_{it} from permanent v_{it} income shocks, which we assume throughout unless otherwise stated. However, for a consumer who simply observed ϵ_{it}, we would have

$$\eta_{it} = \rho_t(1 - \theta_{t+1})\epsilon_{it} + r\theta_{t+1}\epsilon_{it}/(1 + r), \tag{10}$$

which, by analogy with (9), provides a decomposition of the MA innovation ϵ_{it} into the component representing the new information concerning permanent effects and that representing a transitory innovation to income. The permanent effects component can be thought of as capturing news about both current and *past* permanent effects since

$$E\left(\sum_{j=0} v_{i,t-j} \middle| \epsilon_{it}, \epsilon_{i,t-1}, \ldots\right) - E\left(\sum_{j=0} v_{i,t-j} \middle| \epsilon_{i,t-1}, \ldots\right) = (1 - \theta_{t+1})\epsilon_{it}.$$

The decomposition (10) therefore represents the best prediction of the split between permanent and transitory components given θ_{t+1}.

The connection between consumption and income innovations in (9) can be used to link the growth in the variance of consumption and in the covariance of consumption and income to the variances of the underlying components in the income process. From this expression the identification of the growth in transitory and permanent variances can be related to the growth in consumption and income variances as is shown in the following proposition.

PROPOSITION 3. For individuals in a cohort k,

$$\Delta \operatorname{var}_{kt}(c) = \frac{1}{\rho_t^2} \frac{r^2}{(1 + r)^2} \operatorname{var}_{kt}(u) + \operatorname{var}_{kt}(v); \tag{11}$$

$$\Delta \operatorname{cov}_{kt}(c, y) = \Delta \left[\frac{1}{\rho_t} \frac{r}{(1 + r)} \operatorname{var}_{kt}(u)\right] + \operatorname{var}_{kt}(v); \tag{12}$$

8. $\eta_{it} = (r/(1 + r)) \; \Sigma_{k=0}^{T-t} (1 + r)^{-k}(E_t - E_{t-1})y_{t+k}$—see Deaton and Paxson [1994], for example.

$$(13)\quad \Delta\,\mathrm{var}_{kt}(y) - \Delta\,\mathrm{var}_{kt}(c) = \left[1 - \frac{1}{\rho_t^2}\frac{r^2}{(1+r)^2}\right]\mathrm{var}_{kt}(u) - \mathrm{var}_{kt-1}(u).$$

Proof. Equations (11) and (12) follow from (8) and (9). Substituting for $\mathrm{var}_{kt}(v)$ in (7) from (11) gives (13).

Intuitively, the growths in the variance of consumption and in the covariance of consumption and income are dominated by permanent inequality. The proposition shows that the difference of differences in the variances eliminates the variance of the permanent shocks. For large $T - t$ and small r, these results take a particularly simple form.

COROLLARY. For $T - t$ large and r small,

$$(14)\quad \Delta\,\mathrm{var}_{kt}(c) \simeq \mathrm{var}_{kt}(v);$$

$$(15)\quad \Delta\,\mathrm{cov}_{kt}(c, y) \simeq \mathrm{var}_{kt}(v);$$

$$(16)\quad \Delta\,\mathrm{var}_{kt}(y) - \Delta\,\mathrm{var}_{kt}(c) \simeq \Delta\,\mathrm{var}_{kt}(u).$$

Removing the growth in the consumption variance from the growth in the income variance eliminates the permanent inequality term. A higher growth in income variance than in consumption variance must imply a rise in the variance of transitory shocks. In general, the difference provides a lower bound on the growth in transitory uncertainty but for large $T - t$ and small r the corollary shows that growth in short-run income uncertainty is exactly measured by the difference in growth between income inequality and in consumption inequality. To measure the growth in the variance of the transitory shocks to income, we therefore suggest the use of a difference-of-differences estimator (16).

Moreover, for large $T - t$ and small r, individuals consume their permanent income, so that $\Delta\,\mathrm{var}_{kt}(c)$ and $\Delta\,\mathrm{cov}_{kt}(c, y)$ each equal the variance of the permanent shocks. A rise in the variance of permanent income shocks would be reflected in an *acceleration* in their growth. If we have data on both, then (14) and (15) provide one overidentifying restriction per period to use in improving precision of estimates.

Liquidity constraints, by exaggerating the effect of transitory shocks on consumption growth for some consumers, can be thought of in the same way as the finite T model. It remains true that changes in the variance of incomes that are not also reflected in changes in the variance of expenditures can be attributed to transitory shocks. In a simple model in which a fixed proportion of consumers is constrained to consume their income, the difference in the growth of the variance in consumption and income now only identifies the growth in the variance of transitory shocks to the unconstrained group.

Even if u_{it} and v_{it} are not distinguished by the consumer, who observes only ϵ_{it} in (6), then

$$\Delta\ \mathrm{var}_{kt}\,(y) - \Delta\ \mathrm{var}_{kt}\,(c) \simeq \Delta\left(\frac{1-(1-\theta_{t+1})^2}{\theta_{t+1}}\right)\mathrm{var}_{kt}\,(u).$$

A similar path for both variances still suggests a stable pattern of short-run income uncertainty over time.

C. Prudent Consumers and Precautionary Saving

Proposition 3 extends naturally to preferences that admit precautionary saving. An analogous relationship between the variances of consumption and income can be seen to hold under Constant Absolute Risk Aversion (CARA) preferences. Caballero [1990, p. 128] has shown that for such preferences and for income processes of the sort considered here with nonconstant variances,

$$\Delta c_{it} = \Gamma_{kt} + \zeta_{kt} + (1/\rho_{kt})\eta_{it}, \tag{17}$$

where $(1/\rho_{kt})\eta_{it}$ is the annuitized income innovation as previously, Γ_{kt} is the slope of consumption paths within the cohort, and ζ_{kt} is a term taking into account revisions to variance forecasts. With normally distributed innovations, Γ_{kt} is proportional to the expected variance of next period's consumption innovations.[9] Since both Γ_{kt} and ζ_{kt} are assumed constant within cohorts, the results of Proposition 3 continue to hold in this setting. It is important therefore to choose cohort groups which, wherever possible, are sufficiently homogeneous for these constancy assumptions to be valid.[10]

To extend to the case of decreasing absolute risk aversion, we consider Constant Relative Risk Aversion preferences. Assume, as above, that the discount rate equals the real interest rate, and

9. $\Gamma_{kt} = (1/\beta)\ \ln\ E_t e^{[-\beta(1/\rho_{kt})\eta_{it}+\zeta_{kt}]}$, where β is the coefficient of absolute risk aversion.

10. If individuals within a cohort group k are subject to different changes in the variance of permanent shocks, divergences in the growth in variances of consumption and income could not necessarily be given the simple interpretation we establish here.

suppose that a stochastic process similar to that described above applies to the logarithm of income:

$$\ln y_{it} = \ln y_{it}^{p} + u_{it} \qquad \text{for } i \in k$$
$$\ln y_{it}^{p} = \ln y_{it-1}^{p} + v_{it},$$

where u_{it} and v_{it} are assumed orthogonal with properties as above. In such a case it is possible to approximate[11] the growth of individual consumption by an expression similar to (17):

$$\Delta \ln c_{it} \simeq G_{kt} + z_{kt} + \ln\,[L_{it}/(E_{t-1}L_{it})], \tag{18}$$

where G_{kt} and z_{kt} have similar interpretations as the slope of consumption paths within the cohort and as a term taking into account revisions to variance forecasts. Here L_{it} denotes the value of current financial wealth plus the present value of future earnings. If $(1/E_tL_{it})\ \Sigma_{k=0}^{T-t}\ (1 + r)^{-k}E_{t-1}y_{it+k} \simeq 1$, i.e., L_{it} consists mainly of the present value of future earnings, as would seem reasonable for younger cohorts, then[12]

$$\Delta \ln c_{it} \simeq G_{kt} + z_{kt} + v_{it} + ru_{it}/\rho_t(1 + r) \tag{19}$$

which is analogous to the consumption growth process in (8), and the decomposition results of Proposition 3 and the Corollary apply. However, they now refer to identification of the growth in variances of *proportionate* transitory and permanent shocks.

D. Alternative Income Processes

In this subsection we investigate the robustness of these results to modifications to the income process such as allowance for common shocks, deterministic trends, and cross-sectional correlation between the distributions of shocks and of past incomes.

Common Shocks. Common shocks to the income process that impact in the same way across all individuals in the cohort will be captured by the Γ_{kt} term in (17) and will have no impact on within-cohort consumption or income inequality. However, we may wish to allow for common effects that are distributed unevenly across individuals. For example,[13] suppose that we

11. The proof of the approximation is in Appendix 1 and has some similarity to that in Skinner [1988, p. 252], though Skinner establishes his result only for a case in which all shocks are transitory and normally distributed with constant variance. See also Blundell and Stoker [1995].
12. For proof of this further step refer again to Appendix 1.
13. See, for instance, the use of this process in Deaton [1992, p. 148].

define e_{kt} to be a common transitory shock to incomes in cohort k at time t, and suppose that it is distributed across individuals in the cohort according to an individual parameter γ_i. The income process (3) then becomes

$$y_{it} = y_{it}^p + u_{it} + \gamma_i e_{kt} \qquad \text{for } i \in k, \tag{20}$$

where u_{it} is again the purely idiosyncratic transitory shock in period t. Suppose that the permanent component has an analogous decomposition:

$$y_{it}^p = y_{i,t-1}^p + v_{it} + \delta_i v_{kt}. \tag{21}$$

Assuming that $\text{cov}_{kt}(y_{-1},\gamma) = \text{cov}_{kt}(y_{-1},\delta) = 0$, in this case the decomposition results in Proposition 3 generalize to allow the identification of permanent and transitory components in the sum of both idiosyncratic and common contributions to the growth in income uncertainty.[14]

Cross-Section Correlation between Shocks and Past Incomes. Although, when averaging across time, transitory or permanent shocks will be uncorrelated with historical events, this may not be true in a given cross section. That is, in any particular period the distribution of shocks to income may be related to the position of individuals in the income or consumption distribution. For example, this would occur if the γ_i in income process (20) were correlated in the cross section with $y_{i,t-1}$ even though e_{kt} is uncorrelated with $y_{i,t-1}$ over time. In this case $\text{cov}_{kt}(y_{-1},\gamma e) \neq 0$, but $\lim_{t\to\infty}(1/T)\Sigma_t \text{cov}_{kt}(y_{-1},\gamma e) = 0$.[15]

With income processes of this general type, the covariances $\text{cov}_{kt}(y_{-1},v)$ and $\text{cov}_{kt}(c_{-1},v)$ and the covariances $\text{cov}_{kt}(y_{-1},u)$ and $\text{cov}_{kt}(c_{-1},u)$ would be nonzero. In general, if either permanent or transitory shocks are correlated with past incomes and expenditures, then it is clear that $\Delta\ \text{var}_{kt}(c) \neq \Delta\ \text{cov}_{kt}(c,y)$ and the overidentifying restrictions discussed above will fail to hold. Our assumptions are therefore testable against alternatives of this sort. To evaluate whether it is still possible to use cross-sectional variances and covariances of incomes and expenditures to draw inferences about changes in the variances of the permanent and transitory shocks, it is interesting to consider the following separate cases.

Suppose that transitory shocks are uncorrelated with past

14. For CRRA preferences the impact would be a common proportionate effect across all individuals.

15. See also Attanasio and Jappelli [1997].

incomes and consumptions, $\text{cov}_{kt}\,(y_{-1},u) = \text{cov}_{kt}\,(c_{-1},u) = 0$, but allow for correlations with permanent shocks, so that $\text{cov}_{kt}\,(y_{-1},v)$ and $\text{cov}_{kt}\,(c_{-1},v)$ are both nonzero. The simple difference-of-differences estimator $\Delta\,\text{var}_{kt}\,(c) - \Delta\,\text{var}_{kt}\,(y)$ will no longer serve as an estimate of the change in the variance of transitory shocks. However, writing[16]

$$\Delta\,\text{var}_{kt}\,(c) + \Delta\,\text{var}_{kt}\,(y) - 2\Delta\,\text{cov}_{kt}\,(c,y) = \Delta\,\text{var}_{kt}\,(u), \tag{22}$$

it can be seen that knowledge of the covariance term $\Delta\,\text{cov}_{kt}\,(c,y)$ is sufficient to identify the growth in the variance of transitory shocks. Notice, incidentally, that (22) can also be written in terms of the growth of the variance of deviations of income and consumption as

$$\Delta\,\text{var}_{kt}(y - c) = \Delta\,\text{var}_{kt}\,(u). \tag{23}$$

In contrast, suppose that permanent shocks are uncorrelated with past incomes and consumptions, $\text{cov}_{kt}\,(y_{-1},v) = \text{cov}_{kt}\,(c_{-1},v) = 0$, but allow for correlations with transitory shocks, $\text{cov}_{kt}\,(y_{-1},u)$ and $\text{cov}_{kt}(c_{-1},u)$ are both nonzero. Then $\Delta\,\text{var}_{kt}\,(c)$ still provides an estimate of the variance of permanent shocks, but there is no way, in general, to combine this or $\Delta\,\text{cov}_{kt}\,(c,y)$ with $\Delta\,\text{var}_{kt}\,(y)$ to recover the change in the variance of transitory shocks.

Idiosyncratic Trends. A further generalization results in the incorporation of idiosyncratic deterministic trends ϕ_i in the income processes:[17]

$$y_{it} = y_{it-1} + u_{it} - u_{it-1} + v_{it} + \phi_i. \tag{24}$$

Then, for large $T - t$ and small r,

$$\Delta\,\text{var}_{kt}\,(y) = \Delta\,\text{var}_{kt}\,(u) + \text{var}_{kt}\,(v) + \text{var}_{k}\,(\phi) + 2\,\text{cov}_{kt-1}\,(y,\phi),$$

$$\Delta\,\text{var}_{kt}\,(c) \simeq \text{var}_{kt}\,(v).$$

Thus,

$$\Delta\,\text{var}_{kt}\,(y) - \Delta\,\text{var}_{kt}\,(c) \simeq \Delta\,\text{var}_{kt}\,(u) + \text{var}_{k}\,(\phi) + 2\,\text{cov}_{kt-1}\,(y,\phi),$$

and the difference between the growth in income and expenditure variances fails to identify the growth in the transitory variance.

The presence of idiosyncratic deterministic trends would also undermine the testable moment restriction (14) on equality of

16. See Blundell and Preston [1997] for the derivation.

17. We could derive this by supposing permanent income to be the sum of a stochastic component y_{it}^p as before and a deterministic component $y_{it}^d = y_{it-1}^d + \phi_i$.

growth in the consumption variance and consumption-income covariance since

$$\Delta \operatorname{cov}_{kt}(c,y) \simeq \operatorname{var}_{kt}(v) + 2\operatorname{cov}_{kt}(c,\phi) \neq \Delta \operatorname{var}_{kt}(c).$$

Thus, unless such trends are absent, the growth in the expenditure variance will not equal the growth in the expenditure-income covariance.[18] Our assumptions are therefore also testable against alternatives of this sort.

IV. The Growth in Short-Term Income Risk

A. Data

The data used in this study are drawn from the 1968–1992 British Family Expenditure Survey (FES) and are described in the Data Appendix. The long time series of household level data on consumption and income available in Britain and the rapid change in income inequality over this period make it an ideal base for this analysis.

The measure of income used in this paper is current weekly net income of the household, and the measure of expenditure is total current weekly spending by the household on all goods excluding durables but including housing. Expenditure and income of different family types are adjusted onto a comparable basis using equivalence scales that take account of numbers of adults and children in different age ranges, the precise form of the scale being given in the Data Appendix. There are clearly advantages and disadvantages to excluding durables from our measure of consumption. Durable expenditure does not measure consumption of service flows, and this discrepancy is likely to vary systematically over the cycle. However, by excluding durables, we are assuming that there is little substitution between durables and nondurables over time and with age. Although our preference is to exclude durable expenditures, given the potential sensitivity of our conclusions to this, we have repeated all our analysis including durable expenditures and briefly comment on specific comparisons when we discuss the results below. In summary, the conclusions we draw at both the cohort and aggregate level are unaffected by the inclusion of durable expenditures.[19]

18. Blundell and Preston [1997] consider other similar processes.
19. A full set of comparison figures and tables are available from the authors on request.

As argued above, we base most of our analysis on date-of-birth cohorts to avoid some of the obvious biases that result from inequality across cohorts. We know, for instance, from Deaton and Paxson [1994] that one should expect within-cohort inequality to rise with age (see (14)) and also that the age composition of our population is changing. The cohorts are defined by ten-year bands for age of birth of head of household. This enables us to calculate measures of inequality in each year for groups of fairly similar age. Sample sizes for each cohort are given in Appendix 2.

B. Results

For each cohort we calculate the sample variances and covariances of income and expenditure in each year and the associated variance-covariance matrix of these statistics (see Table I). Figure I presents within-cohort paths of the same statistics for the four central cohorts. (These figures can be interpreted as the outcome of nonparametric regressions on a complete set of interacted year and cohort dummies.)

There is a systematic evolution of the within-cohort variances over the whole 25-year period considered. Figure I shows that within-cohort income and expenditure variances increase with age as predicted. However, there is little evidence of the sort of acceleration in expenditure inequality that would be indicative of increasing within-cohort variance of permanent shocks. Nonetheless, income variance rises faster than expenditure variance in the latter half of the 1980s. According to (16), this points to evidence of rising short-term uncertainty. These within-cohort results are robust to the inclusion of durable expenditures in our definition of expenditure. Similar results also come through if we use only those who have stayed in education beyond the compulsory school leaving age.

Two sets of estimates were calculated for the changes in the variances of the underlying shocks. First, we estimate $\Delta\ \mathrm{var}_{kt}(u)$ by the difference between the first differences of the sample income and expenditure variances (as suggested by (16)) and $\Delta\ \mathrm{var}_{kt}(v)$ by the second difference of the sample expenditure variance. A second set of estimates also makes use of the information in the sample covariances, imposing the full set of moment restrictions in (14), (15), and (16) by minimum distance estimation (MDE). These imply one overidentifying moment restriction per period, and the method provides a χ^2 test of these overidentifying restrictions (which can be interpreted as a test for the

TABLE I
VARIANCES AND COVARIANCES

Year	$\mathrm{var}_{kt}(c)$	Born 1920s $\mathrm{var}_{kt}(y)$	$\mathrm{cov}_{kt}(c,y)$	$\mathrm{var}_{kt}(c)$	Born 1930s $\mathrm{var}_{kt}(y)$	$\mathrm{cov}_{kt}(c,y)$
1968	0.1064	0.1033	0.0607	0.0960	0.0934	0.0546
1969	0.1145	0.1143	0.0693	0.1126	0.1069	0.0662
1970	0.1188	0.1119	0.0701	0.1109	0.1080	0.0641
1971	0.1143	0.1179	0.0676	0.1127	0.1272	0.0742
1972	0.1199	0.1222	0.0717	0.1164	0.1188	0.0690
1973	0.1231	0.1179	0.0688	0.1209	0.1021	0.0636
1974	0.1196	0.1112	0.0627	0.1116	0.1106	0.0647
1975	0.1194	0.1090	0.0599	0.1266	0.1101	0.0695
1976	0.1285	0.1130	0.0631	0.1191	0.1065	0.0609
1977	0.1156	0.1029	0.0576	0.1167	0.1042	0.0659
1978	0.1299	0.1201	0.0653	0.1217	0.1120	0.0667
1979	0.1479	0.1279	0.0691	0.1259	0.1202	0.0700
1980	0.1385	0.1469	0.0689	0.1192	0.1259	0.0719
1981	0.1324	0.1882	0.0781	0.1293	0.1436	0.0799
1982	0.1282	0.1729	0.0657	0.1345	0.1524	0.0869
1983	0.1420	0.2015	0.0668	0.1444	0.1724	0.0960
1984	0.1349	0.1799	0.0620	0.1578	0.1605	0.0934
1985	0.1392	0.2544	0.0668	0.1445	0.1896	0.0919
1986				0.1521	0.1923	0.0943
1987				0.1737	0.2485	0.1055
1988				0.1725	0.2401	0.1057
1989				0.1695	0.3109	0.1035
1990				0.1811	0.3047	0.1220
1991				0.1725	0.3619	0.1132
1992				0.1621	0.4784	0.0769

Year	$\mathrm{var}_{kt}(c)$	Born 1940s $\mathrm{var}_{kt}(y)$	$\mathrm{cov}_{kt}(c,y)$	$\mathrm{var}_{kt}(c)$	Born 1950s $\mathrm{var}_{kt}(y)$	$\mathrm{cov}_{kt}(c,y)$
1968	0.1093	0.1065	0.0519			
1969	0.1038	0.1145	0.0588			
1970	0.1071	0.1168	0.0632			
1971	0.1147	0.1326	0.0698			
1972	0.1184	0.1374	0.0759			
1973	0.1205	0.1213	0.0696			
1974	0.1181	0.1203	0.0685			
1975	0.1199	0.1237	0.0704	0.1199	0.1324	0.0686
1976	0.1234	0.1219	0.0714	0.1132	0.1394	0.0682
1977	0.1217	0.1190	0.0735	0.1233	0.1349	0.0729
1978	0.1223	0.1229	0.0756	0.1261	0.1297	0.0800
1979	0.1384	0.1247	0.0794	0.1383	0.1446	0.0808
1980	0.1344	0.1323	0.0810	0.1380	0.1715	0.0922
1981	0.1338	0.1568	0.0883	0.1488	0.1727	0.1028
1982	0.1430	0.1514	0.0940	0.1428	0.1620	0.0999
1983	0.1441	0.1637	0.0971	0.1557	0.1596	0.1033
1984	0.1538	0.1557	0.1016	0.1522	0.1852	0.1134
1985	0.1544	0.1668	0.1064	0.1824	0.1950	0.1315
1986	0.1823	0.1759	0.1092	0.1737	0.1991	0.1227
1987	0.1787	0.2088	0.1261	0.1938	0.2297	0.1461
1988	0.1824	0.2101	0.1222	0.1916	0.2263	0.1492
1989	0.1734	0.2430	0.1263	0.1883	0.2655	0.1468
1990	0.1947	0.2503	0.1438	0.2139	0.2790	0.1698
1991	0.1767	0.2928	0.1251	0.2120	0.2872	0.1813
1992	0.1801	0.2836	0.1300	0.2004	0.3337	0.1658

This table presents variances of log expenditure and log income per equivalent adult and their covariances within cohorts.

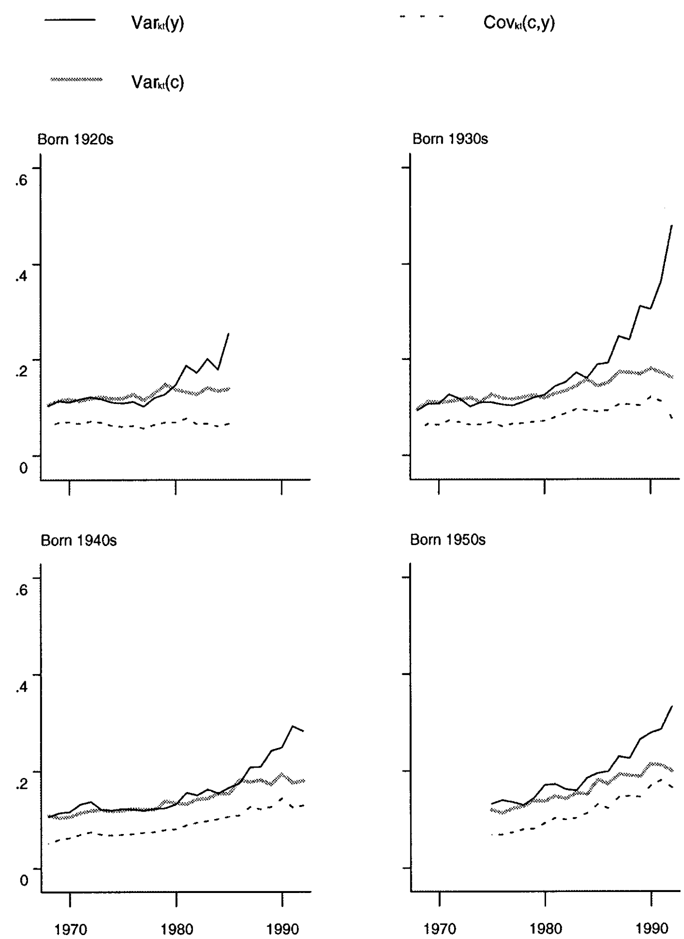

FIGURE I
Variances of Expenditures and Income and Their Covariance by Cohort, 1968–1992

specification of the income process). These procedures were repeated using the logarithms of income and expenditure. We concentrate in the body of the paper on the results using logs but also report the results using levels in Appendix 6.

The estimates in Table II use the difference-of-differences result in equation (16) of the corollary to calculate the growth in the transitory variance $\Delta\ \mathrm{var}_{kt}(u)$ from the difference in rates of growth of consumption and income variances, and these results

TABLE II
ESTIMATES OF CHANGES IN THE VARIANCES OF TRANSITORY SHOCKS TO INCOME

Estimated based on variances alone				
Year	Born 1920s $\Delta\ \text{var}_{kt}(u)$	Born 1930s $\Delta\ \text{var}_{kt}(u)$	Born 1940s $\Delta\ \text{var}_{kt}(u)$	Born 1950s $\Delta\ \text{var}_{kt}(u)$
1969–1976	−0.0153	−0.0070	−0.0121	
	(0.0077)	(0.0075)	(0.0082)	
1976–1984	−0.0605	0.0154	0.0034	0.0068
	(0.0105)	(0.0099)	(0.0093)	(0.109)
1984–1992		0.3135	0.1015	0.1002
		(0.0623)	(0.0201)	(0.0273)
Minimum distance estimates				
Year	Born 1920s $\Delta\ \text{var}_{kt}(u)$	Born 1930s $\Delta\ \text{var}_{kt}(u)$	Born 1940s $\Delta\ \text{var}_{kt}(u)$	Born 1950s $\Delta\ \text{var}_{kt}(u)$
1969–1976	0.0036	0.0038	−0.0064	
	(0.0054)	(0.0053)	(0.0061)	
1976–1984	0.0641	0.0213	0.0042	−0.0021
	(0.0088)	(0.0074)	(0.0065)	(0.0082)
1984–1992		0.2949	0.0991	0.0967
		(0.0620)	(0.0172)	(0.0259)

Estimates based on variances alone are calculated from (14) using log expenditure and log income per equivalent adult. Minimum distance estimates are calculated from (14), (15), and (16) using the same data. Associated χ^2 tests are presented in Table IV. Full year-by-year estimates are given in the Appendix.

are illustrated in Figure II.[20] For all cohorts there is strong evidence of growth in transitory income variance, particularly in the late 1980s where the 95 percent confidence bands for these changes lie well above zero. This provides strong support for the evidence from the income panel data study by Moffitt and Gottschalk [1995] based on United States Panel Study of Income Dynamics data. Table III contains estimates of the growth in the permanent variance $\Delta\ \text{var}_{kt}(v)$ based on the acceleration in the expenditure variance, and these are illustrated in Figure III. It is clear that none of these changes are significantly different from zero.

Using Covariance Information. Since the Family Expenditure Survey contains detailed information on income *and* expenditure

20. For clarity, Tables II and III present estimates of total changes over three subperiods. Year-by-year changes are presented in Appendices 4 and 5. Figures II and III present year-by-year changes smoothed by taking five-year moving averages.

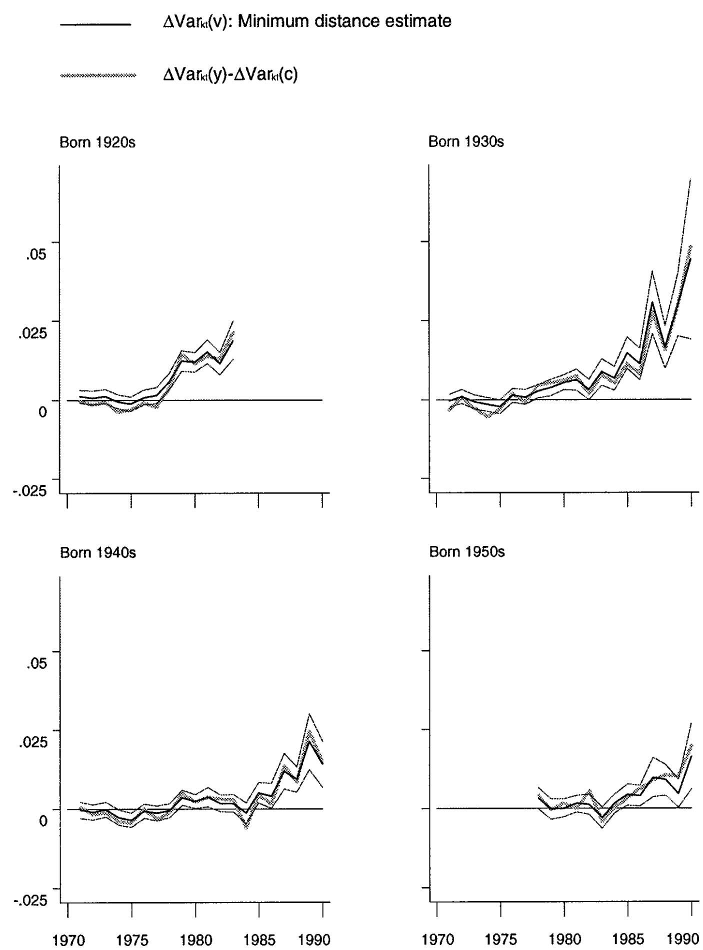

FIGURE II
Estimated Changes in the Variance of Transitory Income Shocks by Cohort, 1968–1992

for each household, we can use covariances to test the robustness of our results. Tables II and III also provide estimates using minimum distance estimation applied to the moment conditions (14), (15), and (16).[21] These estimates are also represented in

21. These are calculated by minimum distance estimation with the asymptotically optimal weights. Conscious of the arguments of Altonji and Segal [1996] for preferring equally weighted minimum distance estimates in small samples, we

TABLE III
ESTIMATES OF CHANGES IN THE VARIANCES OF PERMANENT SHOCKS TO INCOME

Estimates based on variances alone

Year	Born 1920s $\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1930s $\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1940s $\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1950s $\Delta\,\mathrm{var}_{kt}\,(v)$
1969–1976	0.0010	−0.0241	0.0090	
	(0.0090)	(0.0093)	(0.0103)	
1976–1984	−0.0162	0.0208	0.0062	0.0033
	(0.0106)	(0.0114)	(0.0107)	(0.0127)
1984–1992		−0.0238	−0.0063	−0.0080
		(0.0137)	(0.0129)	(0.0133)

Minimum distance estimates

Year	Born 1920s $\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1930s $\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1940s $\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1950s $\Delta\,\mathrm{var}_{kt}\,(v)$
1969–1976	−0.0040	−0.0215	0.0000	
	(0.0074)	(0.0074)	(0.0083)	
1976–1984	−0.0115	0.0116	0.0045	0.0093
	(0.0087)	(0.0095)	(0.0087)	(0.0108)
1984–1992		−0.0201	−0.0023	−0.0208
		(0.0122)	(0.0116)	(0.0123)

Estimates based on variances alone are calculated from (16) using log expenditure and log income per equivalent adult. Minimum distance estimates are calculated from (14), (15), and (16) using the same data. Associated χ^2 tests are presented in Table IV. Full year-by-year estimates are given in the Appendix.

Figures II and III where they are presented with confidence bands.[22] The evidence of growing within-cohort variance of transitory shocks and static within-cohort variance of permanent shocks is strengthened, at least for the two younger cohorts, by the use of covariance information. Although we document an increase in permanent inequality over time for each cohort, there is no evidence of an increase in the variance of the permanent component of income shocks. The growth in permanent inequality within each cohort comes purely from the accumulation of permanent shocks in the consumption growth equation (19).

Note from Table IV that the overidentifying restrictions based on (14), (15), and (16) in the Corollary are acceptable for the

also recalculated results (available on request) using equal weights but found very little difference in results.

22. These estimates are also smoothed in the figures by taking five-year moving averages. Confidence bands are adjusted accordingly.

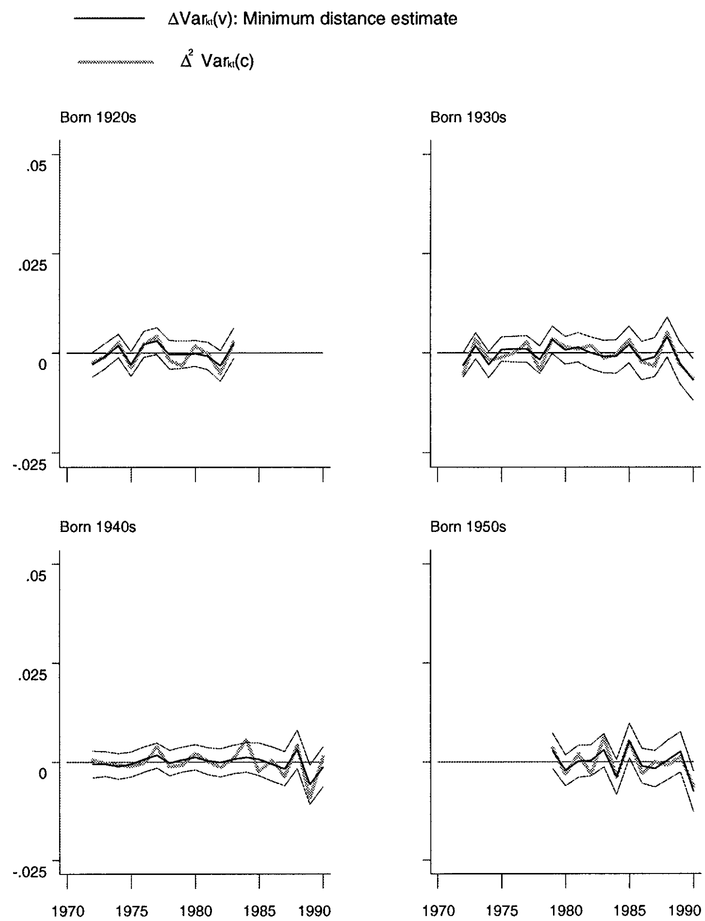

FIGURE III
Estimated Changes in the Variance of Permanent Income Shocks by Cohort, 1968–1992

younger two cohorts only—in other words, exactly those for whom $T - t$ is largest.[23] In principle, for older cohorts it would be possible to estimate the permanent and transitory variances by minimum distance by choosing an appropriate value for r and fixing values for $r/\rho_t(1 + r)$ in (11) to (13).

23. These overidentification tests fail when we include durable expenditures in our expenditure definition, although only marginally for the youngest cohort, and the test statistics remain largest for the older cohorts.

TABLE IV
TESTS OF OVERIDENTIFYING RESTRICTIONS

Born 1920s	Born 1930s	Born 1940s	Born 1950s
$\chi^2_{17} = 81.05$	$\chi^2_{24} = 65.38$	$\chi^2_{24} = 34.46$	$\chi^2_{17} = 27.55$
$P = 0.000$	$P = 0.000$	$P = 0.077$	$P = 0.051$

These are χ^2 tests of the overidentifying moment restrictions in (14), (15), and (16) using log expenditure and log income per equivalent adult.

Cross-Sectional Correlation between Shocks and Past Incomes. It was shown above that neither the difference-in-differences nor the MDE estimate of the change in the variance of transitory shocks would be robust to cross-sectional correlation between permanent shocks and past incomes and expenditures. However, an alternative robust estimate using variance and covariance information was proposed in (22). This estimate is presented, together with the simpler difference-in-differences estimate, in Figure IV. It is evident that the picture of increasing transitory uncertainty is preserved even if one allows for permanent shocks correlated with past incomes with estimated growth in the transitory variance consistently positive throughout the 1980s using either estimate.

The Overall Picture. What is also interesting, however, is the comparison of the growth in inequality within and across cohorts. This can be gauged from Figure V which presents the evolution over time of the variances and covariance of logarithms[24] for the whole sample.[25] There is no a priori reason why the consumption rule (19) itself should produce such a picture even though for each cohort there is growth of consumption inequality. The growth of consumption inequality in Figure V reflects one or more of three possible explanations. First, it may reflect the aging population. Second, it may reflect a higher level of permanent inequality for younger cohorts when comparison is made for the same age. Third, it may reflect a growth in within-cohort variance of permanent components. The later explanation is ruled out from our discussion of Figure III. The first could explain part of the

24. It should be noted that the variance of logarithms is not a wholly satisfactory index of inequality since progressive transfers of income or expenditure can cause it to increase. Results using the variance of levels give a similar picture and are available from the authors on request.

25. This picture is unaffected by the inclusion of durable expenditures.

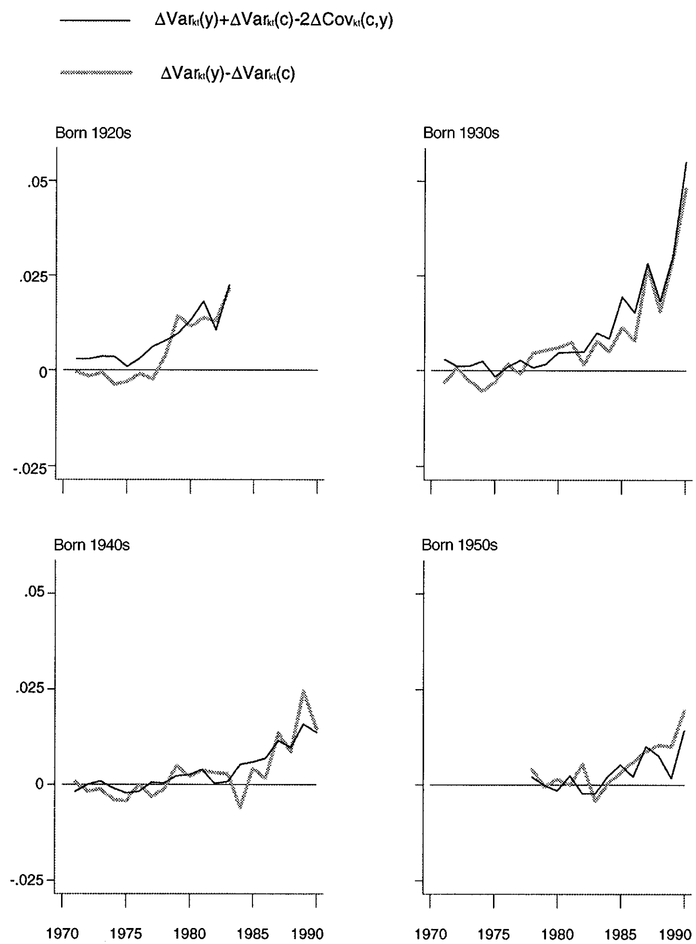

FIGURE IV
Estimated Changes in the Variance of Transitory Income Shocks by Cohort, 1968–1992: Robustness to Permanent Shocks Correlated with Past Incomes

growth in the overall picture, but the second explanation is also supported by the data. To see this, consider Figure VI. Figure VI plots the variances against age for all cohorts in one diagram. The results are striking. Inequality in both income and consumption was sharply greater in the late 1980s for all cohorts than it had been for preceding cohorts at a similar age. We can now see that younger cohorts face not only an increase in the variance of

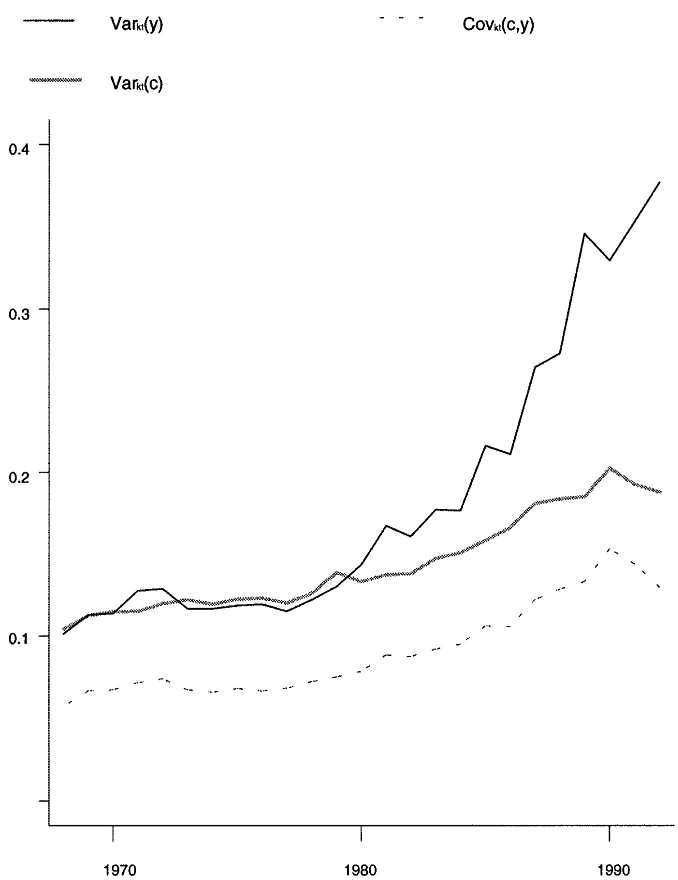

FIGURE V
Variances of Expenditure and Income and Their Covariance for the Whole Sample, 1968–1992

transitory components by comparison with older cohorts when they were at a similar age but that this is also true for permanent income inequality. As the variance of permanent shocks does not appear to have risen for any cohort, the implication is that this difference reflects an increase in initial permanent income inequality for younger cohorts. The explanation for a growth in overall consumption inequality can therefore be attributed in part to an aging population and in part to new cohorts facing higher levels of initial income inequality.

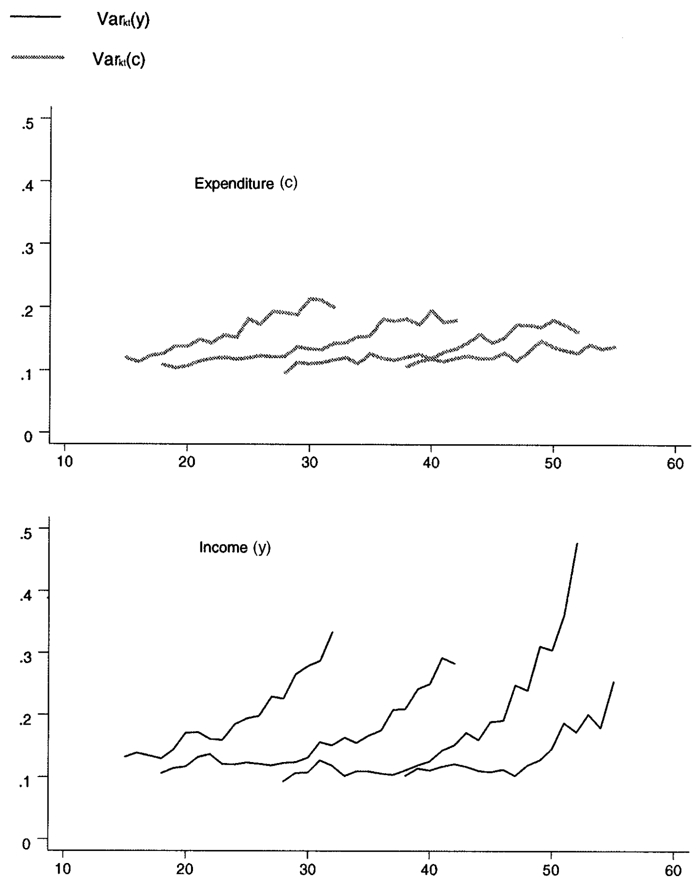

Figure VI
Variances of Expenditures and Income for Each Cohort by Age, 1968–1992

V. Conclusions

This paper has shown how a comparison of the growth paths of income and consumption inequality can together be used to document the growth in short-term income risk. We derive conditions under which the growth of variances and covariances of income and consumption can be used to separately identify the growth in the variance of permanent and transitory income shocks. This sheds new light on the debate concerning the use of consumption versus income in assessing the evolution of inequality.

Comparisons within cohorts are shown to carry more reliable welfare implications than comparisons across cohorts at any point in time. Increasing short-term risk will inflict a welfare cost on risk-averse households, and we derive the set of preferences under which the level of precautionary saving can be used as a money metric measure of the individual welfare loss from income risk.

We use a sample of households from the British Family Expenditure Survey over the period 1968–1992. Over this period both income and consumption inequality is shown to have risen for all cohorts. However, there is striking evidence that income inequality has risen faster than consumption inequality in recent years especially for younger cohorts. By plotting the difference of differences between consumption and income inequality across this period, we are able to measure the precise path of growth in short-term income risk over this period. The paths of consumption-income covariances within cohorts provide corroborative evidence of growing short-term uncertainty.

Data Appendix

The data used here are from the 1968–1992 British Family Expenditure Survey (FES). The FES is an annual survey conducted with the principal purpose of determining the basket of goods used to construct the Retail Price Index. In a typical year the FES contains information on around 7000 households. In general, the households form a representative sample, but excluded are those not living in private houses, such as residents of residential homes.

The measure of income used in this paper is current weekly net income of the household and consists of earnings from main and subsidiary jobs (net of tax and national insurance contributions), net profits from self-employment, all social security benefits received, allowances from nonmembers of the household, benefits from friendly societies, and children's incomes.

Households participating in the FES are asked to complete a diary detailing all their spending.[26] In this paper, expenditure is

26. All of what has been said so far assumes that income and expenditure are accurately recorded in the data. We would not wish to ignore the consideration that expenditure may often be better measured, particularly for groups such as the self-employed. On the other hand, infrequency of purchase may well be important given that diary records are kept over a two-week period (see Kay, Keen, and Morris [1984]).

defined as total current weekly spending by the household on all goods excluding durables but including housing.[27]

Expenditure and income of different family types are adjusted onto a comparable basis, using the equivalence scales based on McClements [1977] and favored by the United Kingdom Department of Social Security, and expressed in 1992 pounds for a childless couple. These equivalence scales depend on number of adults and numbers of children in various age ranges.[28] We remove the households with the highest and lowest 2 percent of incomes and expenditures in each year so as to enhance robustness of the results.

We split the samples into cohorts defined by ten-year bands for age of birth of head of household. Appendix 2 shows a cross tabulation of the numbers of households in each annual FES with the head of household's date of birth falling into particular ranges. We drop from the analysis any such cohort in any year in which it could include households with head aged over the statutory retirement age of 65 or in which the sample size falls below 300. This leads us to concentrate on the four central cohorts with head born between 1920 and 1960. Sample sizes and summary statistics for income and expenditure by year for households belonging to these cohorts are given in Appendix 3.

Sensitivity of results to the use of levels rather than logarithms is shown in Appendix 4. It is clear that the restrictions implied by the theory are more acceptable when using logs though the main conclusions about rising transitory variances remain evident when using levels. Sensitivity to choice of equivalence scales was investigated, and results were found to be qualitatively similar with alternative scales.[29]

Appendix 1: Proof of Propositions

A. Proof of Proposition 1

(i) If $b_i = b_j$, then $\mathbf{p}_i = \mathbf{p}_j$. The result then follows trivially.

(ii) If $b_i \neq b_j$, then it may be that $\mathbf{p}_i \neq \mathbf{p}_j$. Hence $c_t(U,\mathbf{p})$ cannot

27. For a detailed consideration of the issues involved in the treatment of housing costs, see Johnson and Webb [1991].

28. Specifically, the scale assigns a first adult 0.55 times the cost for a couple, second and third adults 0.45, and subsequent adults 0.40. Children have a cost of 0.07 times that of a couple if aged 2 or under, 0.18 if aged 3 or 4, 0.22 if aged 5 to 10, 0.27 if aged 11 to 16, and 0.38 if aged 17 or 18.

29. For details see Blundell and Preston [1997].

depend on $\mathbf{p}$; i.e., $c_t(U,\mathbf{p}) = f_t(U)$. If $e(U,\mathbf{p})$ is the expenditure function then Shephard's lemma requires that

$$\frac{\partial e(U,\mathbf{p})}{\partial p_t} = f_t(U) \Leftrightarrow e(U,\mathbf{p}) = \sum_t p_t f_t(U) + \varphi(U) \tag{25}$$

and $\varphi(U) = 0$ by homogeneity. If $U = \min_t u_t(c_t)$, then

$$e(U,\mathbf{p}) = \min_c \left[\sum_t p_t c_t \middle| \min_t u_t(c_t) \geq U \right]$$
$$= \sum_t p_t f_t(U), \text{ where } u_t(f_t(U) \equiv U$$

as required. Note also that (25) implies that the distance function $d(U,\mathbf{c})$ (see Deaton and Muellbauer [1980, p. 53] takes the form,

$$d(U,\mathbf{c}) = \min_{\mathbf{p}} \left[\sum_t p_t c_t \middle| \sum_t p_t f_t(U) = 1 \right] = \min_r \frac{c_t}{f_t(U)} \Rightarrow U = \min_t u_t(c_t).$$

Hence both necessity and sufficiency are established.

(iii) Since it is possible that $s \neq t$, then $c_t(U,\mathbf{p})$ cannot depend on t; i.e., it must be that $c_t(U,\mathbf{p}) = f(U,\mathbf{p})$. By quasi concavity of preferences this can be so only if $f(U,\mathbf{p}) = f(U)$. It is obvious that the same condition would ensure welfare comparisons across cohorts. Then similar reasoning establishes $U = \min_t u(c_t)$ as the corresponding direct representation for preferences.

B. Proof of Proposition 2

Define $\hat{c}_{it}$ by $u_t(\hat{c}_{it}) \equiv Eu_t(c_{it})$. Then

$$\sum_{t=0} u_t(\tilde{c}_{it}) = u_t(c_{i0}) + \sum_{t=1} u_t(\hat{c}_{it}). \tag{26}$$

Note that

$$\frac{u'_0(\tilde{c}_{i0})}{u'_t(\tilde{c}'_{it})} = \frac{u'_0(c_{i0})}{Eu'_t(c_{it})} = p_{it}. \tag{27}$$

Since $U(.,.)$ is quasi concave and noting (26), $c_{i0} = \tilde{c}_{i0}$ if and only if

$$\frac{u'_0(c_{i0})}{u'_t(\hat{c}_{it})} = p_{it}$$

for all t which is true given (27) if and only if $u'_t(\hat{c}_{it}) = Eu'_t(c_{it})$.

To see that CARA is sufficient for this, note that, given CARA,

$$u'_t(\hat{c}_{it}) = -\beta_t u_t(\hat{c}_{it}) = -\beta_t E u_t(c_{it}) = E u'_t(c_{it}).$$

To establish necessity, consider the case where risk is small. A Taylor expansion around $c_{it} = \hat{c}_{it}$ yields

$$(28) \quad E u'_t(c_{it}) \simeq u'_t(\hat{c}_{it}) + E(c_{it} - \hat{c}_{it}) u''_t(\hat{c}_{it}) + ½ E(c_{it} - \hat{c}_{it})^2 u'''_t(\hat{c}_{it}).$$

Similarly, by another Taylor expansion,

$$E u_t(c_{it}) \simeq u_t(\hat{c}_{it}) + E(c_{it} - \hat{c}_{it}) u'_t(\hat{c}_{it}) + \frac{1}{2} E(c_{it} - \hat{c}_{it})^2 u''_t(\hat{c}_{it}).$$

$$\Rightarrow \frac{E(c_{it} - \tilde{c}_{it})}{E(c_{it} - \hat{c}_{it})^2} \simeq \frac{1}{2} \frac{u''_t(\hat{c}_{it})}{u'_t(\hat{c}_{it})}.$$

Substituting into (28) gives

$$E u'_t(c_{it}) \simeq u'_t(\hat{c}_{it}) + \frac{1}{2} E(c_{it} - \hat{c}_{it})^2 \left[u'''_t(\hat{c}_{it}) - \frac{u''_t(\hat{c}_{it})^2}{u'_t(\hat{c}_{it})} \right].$$

Hence $u'_t(\hat{c}_{it}) = E u'_t(c_{it})$ for small enough risks only if

$$u'''_t(\hat{c}_{it}) / u''_t(\hat{c}_{it}) = u''_t(\hat{c}_{it}) / u'_t(\hat{c}_{it}),$$

which holds for all $\hat{c}_{it}$ iff

$$u_t(c_{it}) = A \exp(B c_{it}) + C$$

for some constants A, B, and C. Without affecting behavior, we can set $C = 0$ and the requirement that $U(.,.)$ be increasing and quasi concave ensures $A < 0$, $B < 0$.

C. Proof of (18) and (19)

Let L_{it} denote the value of current financial wealth plus the present value of future earnings at period t, and let $\xi_{it} = L_{it} - E_{t-1} L_{it}$ denote the innovation to L_{it}. Let preferences be Constant Relative Risk Aversion (CRRA) with γ denoting the Arrow-Pratt measure of CRRA (see Skinner [1988, p. 241]), and let the constant interest rate r equal the consumer's subjective discount rate.

Suppose that optimal consumption c_{it} is approximately proportional to L_{it}:

$$c_{it} \simeq \phi_{it} L_{it}$$

for some ϕ_{it} which could depend on uncertain current and future moments of the income process. We know that this is trivially true for $t = T$ since $c_{iT} = L_{iT}$.

By the Euler equation and intertemporal budget constraint,

$$\begin{aligned} c_{it-1}^{-\gamma} &\simeq E_{t-1}(\phi_{it}L_{it})^{-\gamma} \\ &= E_{t-1}(\phi_{it}\{(L_{it-1} - c_{it-1})(1 + r) + \xi_{it}\})^{-\gamma}. \end{aligned}$$

Taking a Taylor expansion around $\phi_{it} = E_{t-1}\phi_{it} \equiv \overline{\phi}_{it}$ and $\xi_{it} = 0$, i.e., $L_{it} = E_{t-1}L_{it} = (L_{it-1} - c_{it-1})(1 + r) \equiv \overline{L}_{it}$ and assuming ϕ_{it} and ξ_{it} to be independent,

$$\begin{aligned} c_{it-1}^{-\gamma} &\simeq (\overline{\phi}_{it}\overline{L}_{it})^{-\gamma}\left[1 + \frac{\gamma(\gamma + 1)}{2}\left\{\frac{\operatorname{var}(\xi_{it})}{\overline{L}_{it}^2} + \frac{\operatorname{var}(\phi_{it})}{\overline{\phi}_{it}^2}\right\}\right] \\ &\equiv (\overline{\phi}_{it}\overline{L}_{it})^{-\gamma}[1 + K_{it}]. \end{aligned}$$

Substituting from the budget constraint,

$$c_{it-1} \simeq \overline{\phi}_{it}(L_{it-1} - c_{it-1})(1 + r)(1 + K_{it})^{-1/\gamma},$$

and thus,

$$\begin{aligned} c_{it-1} &\simeq \frac{\overline{\phi}_{it}(1 + r)(1 + K_{it})^{-1/\gamma}}{1 + \overline{\phi}_{it}(1 + r)(1 + K_{it})^{-1/\gamma}} L_{it-1} \\ &\equiv \phi_{it-1}L_{it-1}. \end{aligned}$$

Hence the supposed approximate proportionality of consumption is established by induction. Furthermore,

$$\frac{c_{it}}{c_{it-1}} \simeq \frac{\phi_{it}L_{it}}{\overline{\phi}_{it}\overline{L}_{it}}(1 + K_{it})^{1/\gamma},$$

and thus,

$$\begin{aligned} \Delta \ln c_{it} &\simeq \frac{1}{\gamma}\ln(1 + K_{it}) + \ln\left(\frac{\phi_{it}}{\overline{\phi}_{it}}\right) + \ln\left(\frac{L_{it}}{E_{t-1}L_{it}}\right) \\ &\equiv G_{kt} + z_{kt} + \ln\left(\frac{L_{it}}{E_{t-1}L_{it}}\right). \end{aligned}$$

By a series of further approximations,

$$\ln\left(\frac{L_{it}}{E_{t-1}L_{it}}\right) \simeq \frac{L_{it} - E_{t-1}L_{it}}{E_{t-1}L_{it}}$$

$$\simeq \frac{1}{E_{t-1}L_{it}} \sum_{k=0}^{T-t} (1+r)^{-k}(E_t - E_{t-1})y_{it+k}$$

$$\simeq \frac{1}{E_{t-1}L_{it}} \sum_{k=0}^{T-t} (1+r)^{-k}E_{t-1}y_{it+k}(E_t - E_{t-1}) \ln y_{it+k}$$

$$\simeq v_{it} + \frac{r}{\rho_t(1+r)} u_{it}$$

if $(1/E_{t-1}L_{it}) \Sigma_{k=0}^{T-t} (1+r)^{-k} E_{t-1} y_{it+k} \simeq 1$.

APPENDIX 2: COHORT SAMPLE SIZES

Year	Born 1920s	Born 1930s	Born 1940s	Born 1950s
1968	1227	1079	653	**2**
1969	1156	1066	734	**8**
1970	1020	966	805	**26**
1971	1155	1040	980	**82**
1972	1160	1025	1073	**142**
1973	1154	949	1076	**214**
1974	995	920	1033	**266**
1975	1080	1016	1154	407
1976	1091	1006	1105	558
1977	1055	987	1170	631
1978	1107	930	1148	717
1979	979	914	1108	832
1980	1011	954	1167	852
1981	1090	1003	1253	987
1982	1107	988	1226	1084
1983	993	941	1105	1043
1984	1091	926	1098	1081
1985	1009	923	1124	1068
1986	**1001**	891	1091	1116
1987	**1118**	855	1092	1192
1988	**1003**	880	1085	1105
1989	**1083**	882	1050	1089
1990	**998**	819	987	1065
1991	**973**	851	936	1078
1992	**1066**	852	1061	1129

Numbers in boldface represent cohorts where some members may be over 65 or sample size is below 300. These cells are not used in the analysis.

Appendix 3: Summary Statistics

Year	Expenditure mean	Std dev	Income mean	Std dev	Sample size
1968	136.33	47.38	161.76	52.48	2961
1969	142.67	51.13	162.93	55.36	2964
1970	145.94	53.71	167.83	58.50	2817
1971	146.19	53.07	168.65	61.47	3257
1972	153.80	57.45	179.45	64.79	3400
1973	161.50	60.96	186.66	64.28	3393
1974	165.04	61.77	185.97	63.68	3214
1975	160.78	60.11	182.80	62.94	3657
1976	157.02	59.05	177.14	61.18	3760
1977	155.40	57.63	174.24	58.86	3843
1978	162.55	61.89	189.36	65.80	3902
1979	171.97	68.41	195.50	69.28	3833
1980	167.59	65.61	195.47	72.94	3984
1981	168.12	67.55	189.82	75.81	4333
1982	163.05	64.78	183.33	73.06	4405
1983	171.17	71.03	185.34	76.71	4082
1984	173.14	72.55	186.43	76.49	4196
1985	176.81	76.39	192.31	84.90	4124
1986	184.56	82.91	197.02	90.18	4099
1987	188.46	88.78	206.23	104.71	4257
1988	191.43	88.94	212.08	107.86	4073
1989	195.12	90.76	211.35	111.18	4104
1990	198.75	98.51	214.40	121.29	3869
1991	200.06	94.36	212.00	120.02	3838
1992	200.91	93.97	208.74	118.18	4108
Total	170.99	74.73	190.27	84.47	94473

This table presents summary statistics for levels of expenditure and income per equivalent adult.

APPENDIX 4: ESTIMATES OF CHANGES IN THE VARIANCES OF TRANSITORY AND PERMANENT SHOCKS TO INCOME BASED ON LOG EXPENDITURE AND LOG INCOME VARIANCES ALONE

	Born 1920s		Born 1930s		Born 1940s		Born 1950s	
Year	$\Delta\ \mathrm{var}_{kt}\ (u)$	$\Delta\ \mathrm{var}_{kt}\ (v)$	$\Delta\ \mathrm{var}_{kt}\ (u)$	$\Delta\ \mathrm{var}_{kt}\ (v)$	$\Delta\ \mathrm{var}_{kt}\ (u)$	$\Delta\ \mathrm{var}_{kt}\ (v)$	$\Delta\ \mathrm{var}_{kt}\ (u)$	$\Delta\ \mathrm{var}_{kt}\ (v)$
1969	0.0028		−0.0031		0.0135			
	(0.0068)		(0.0069)		(0.0090)			
1970	−0.0067	−0.0039	0.0028	−0.0183	−0.0010	0.0088		
	(0.0074)	(0.0107)	(0.0077)	(0.0114)	(0.0086)	(0.0126)		
1971	0.0106	−0.0088	0.0174	0.0035	0.0082	0.0043		
	(0.0075)	(0.0114)	(0.0079)	(0.0123)	(0.0087)	(0.0120)		
1972	−0.0013	0.0101	−0.0121	0.0020	0.0012	−0.0040		
	(0.0075)	(0.0108)	(0.0080)	(0.0116)	(0.0083)	(0.0121)		
1973	−0.0075	−0.0023	−0.0212	0.0006	−0.0183	−0.0015		
	(0.0080)	(0.0113)	(0.0082)	(0.0122)	(0.0079)	(0.0116)		
1974	−0.0032	−0.0067	0.0178	−0.0137	0.0015	−0.0046		
	(0.0080)	(0.0118)	(0.0086)	(0.0129)	(0.0079)	(0.0117)		
1975	−0.0020	0.0033	−0.0155	0.0243	0.0016	0.0042		
	(0.0077)	(0.0118)	(0.0084)	(0.0130)	(0.0078)	(0.0118)		
1976	−0.0052	0.0094	0.0039	−0.0225	−0.0053	0.0017	0.0138	
	(0.0078)	(0.0115)	(0.0078)	(0.0126)	(0.0077)	(0.0117)	(0.0124)	
1977	0.0028	−0.0220	0.0001	0.0050	−0.0012	−0.0052	−0.0146	0.0170
	(0.0079)	(0.0119)	(0.0076)	(0.0119)	(0.0074)	(0.0120)	(0.0106)	(0.0159)
1978	0.0030	0.0271	0.0029	0.0074	0.0033	0.0024	−0.0080	−0.0074
	(0.0079)	(0.0117)	(0.0083)	(0.0121)	(0.0071)	(0.0116)	(0.0097)	(0.0152)
1979	−0.0103	0.0038	0.0039	−0.0007	−0.0143	0.0155	0.0027	0.0094
	(0.0092)	(0.0128)	(0.0089)	(0.0131)	(0.0080)	(0.0120)	(0.0096)	(0.0144)
1980	0.0284	−0.0274	0.0124	−0.0109	0.0115	−0.0201	0.0272	−0.0125
	(0.0102)	(0.0143)	(0.0089)	(0.0132)	(0.0086)	(0.0130)	(0.0106)	(0.0150)
1981	0.0475	0.0032	0.0076	0.0167	0.0251	0.0034	−0.0096	0.0111
	(0.0109)	(0.0139)	(0.0095)	(0.0130)	(0.0087)	(0.0131)	(0.0108)	(0.0151)
1982	−0.0111	0.0019	0.0036	−0.0049	−0.0146	0.0098	−0.0047	−0.0168
	(0.0108)	(0.0130)	(0.0101)	(0.0133)	(0.0088)	(0.0125)	(0.0099)	(0.0146)
1983	0.0147	0.0181	0.0101	0.0047	0.0112	−0.0081	−0.0153	0.0189
	(0.0123)	(0.0129)	(0.0112)	(0.0138)	(0.0096)	(0.0128)	(0.0095)	(0.0136)
1984	−0.0145	−0.0210	−0.0253	0.0035	−0.0177	0.0086	0.0292	−0.0165
	(0.0132)	(0.0140)	(0.0120)	(0.0149)	(0.0104)	(0.0137)	(0.0106)	(0.0142)
1985	0.0702	0.0114	0.0423	−0.0266	0.0105	−0.0091	−0.0204	0.0337
	(0.0165)	(0.0142)	(0.0141)	(0.0158)	(0.0107)	(0.0145)	(0.0116)	(0.0145)
1986			−0.0049	0.0208	−0.0188	0.0273	0.0127	−0.0389
			(0.0150)	(0.0157)	(0.0114)	(0.0153)	(0.0118)	(0.0157)
1987			0.0346	0.0140	0.0365	−0.0316	0.0105	0.0288
			(0.0164)	(0.0172)	(0.0123)	(0.0168)	(0.0120)	(0.0161)
1988			−0.0073	−0.0227	−0.0024	0.0074	−0.0012	−0.0224
			(0.0179)	(0.0189)	(0.0133)	(0.0170)	(0.0122)	(0.0167)
1989			0.0739	−0.0019	0.0419	−0.0128	0.0425	−0.0010
			(0.0301)	(0.0183)	(0.0195)	(0.0170)	(0.0178)	(0.0168)
1990			−0.0177	0.0146	−0.0140	0.0303	−0.0121	0.0289
			(0.0315)	(0.0185)	(0.0203)	(0.0170)	(0.0198)	(0.0175)
1991			0.0657	−0.0201	0.0605	−0.0393	0.0101	−0.0275
			(0.0287)	(0.0202)	(0.0251)	(0.0185)	(0.0168)	(0.0188)
1992			0.1269	−0.0019	−0.0126	0.0214	0.0581	−0.0096
			(0.0663)	(0.0189)	(0.0293)	(0.0176)	(0.0287)	(0.0186)

Estimates are calculated from (14) and (16) using log expenditure and log income per equivalent adult.

Appendix 5: Minimum Distance Estimates of Changes in the Variances of Transitory and Permanent Shocks to Income Using Logs

Year	Born 1920s $\Delta\,\mathrm{var}_{kt}\,(u)$	$\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1930s $\Delta\,\mathrm{var}_{kt}\,(u)$	$\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1940s $\Delta\,\mathrm{var}_{kt}\,(u)$	$\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1950s $\Delta\,\mathrm{var}_{kt}\,(u)$	$\Delta\,\mathrm{var}_{kt}\,(v)$
1969	0.0012		0.0007		0.0037			
	(0.0048)		(0.0050)		(0.0068)			
1970	−0.0030	−0.0068	0.0034	−0.0156	−0.0031	0.0027		
	(0.0052)	(0.0089)	(0.0054)	(0.0089)	(0.0063)	(0.0107)		
1971	0.0094	−0.0063	0.0085	0.0086	0.0099	0.0028		
	(0.0054)	(0.0095)	(0.0058)	(0.0096)	(0.0064)	(0.0101)		
1972	−0.0003	0.0096	−0.0029	−0.0084	−0.0008	−0.0028		
	(0.0056)	(0.0091)	(0.0060)	(0.0096)	(0.0064)	(0.0101)		
1973	−0.0020	−0.0060	−0.0116	−0.0014	−0.0114	−0.0065		
	(0.0057)	(0.0093)	(0.0057)	(0.0097)	(0.0059)	(0.0100)		
1974	−0.0009	−0.0045	0.0075	0.0020	0.0003	0.0006		
	(0.0058)	(0.0097)	(0.0056)	(0.0096)	(0.0059)	(0.0095)		
1975	0.0000	0.0035	−0.0048	0.0083	0.0015	0.0034		
	(0.0057)	(0.0099)	(0.0054)	(0.0098)	(0.0057)	(0.0095)		
1976	0.0004	0.0065	0.0038	−0.0149	−0.0029	−0.0003	0.0074	
	(0.0056)	(0.0096)	(0.0052)	(0.0099)	(0.0056)	(0.0094)	(0.0089)	
1977	−0.0038	−0.0128	−0.0065	0.0103	−0.0048	−0.0005	−0.0093	0.0101
	(0.0054)	(0.0094)	(0.0050)	(0.0094)	(0.0053)	(0.0093)	(0.0079)	(0.0137)
1978	0.0083	0.0183	0.0069	−0.0006	0.0023	0.0004	−0.0124	−0.0011
	(0.0055)	(0.0092)	(0.0053)	(0.0094)	(0.0052)	(0.0093)	(0.0070)	(0.0129)
1979	0.0022	−0.0008	0.0046	0.0018	−0.0026	0.0056	0.0138	−0.0014
	(0.0063)	(0.0103)	(0.0061)	(0.0102)	(0.0056)	(0.0097)	(0.0072)	(0.0127)
1980	0.0217	−0.0134	0.0043	−0.0053	0.0058	−0.0069	0.0158	0.0035
	(0.0070)	(0.0115)	(0.0062)	(0.0108)	(0.0057)	(0.0102)	(0.0078)	(0.0127)
1981	0.0332	0.0066	0.0096	0.0104	0.0170	0.0045	−0.0092	0.0028
	(0.0086)	(0.0114)	(0.0068)	(0.0111)	(0.0061)	(0.0103)	(0.0075)	(0.0128)
1982	−0.0054	−0.0112	0.0015	−0.0027	−0.0108	0.0022	−0.0077	−0.0146
	(0.0094)	(0.0111)	(0.0073)	(0.0112)	(0.0064)	(0.0106)	(0.0070)	(0.0123)
1983	0.0251	0.0153	0.0117	0.0029	0.0095	−0.0046	−0.0052	0.0102
	(0.0108)	(0.0108)	(0.0090)	(0.0118)	(0.0075)	(0.0108)	(0.0068)	(0.0115)
1984	−0.0172	−0.0135	−0.0109	−0.0053	−0.0124	0.0038	0.0121	−0.0002
	(0.0118)	(0.0116)	(0.0098)	(0.0130)	(0.0079)	(0.0115)	(0.0074)	(0.0118)
1985	0.0586	0.0147	0.0316	−0.0099	0.0055	−0.0024	−0.0052	0.0163
	(0.0161)	(0.0118)	(0.0117)	(0.0134)	(0.0074)	(0.0119)	(0.0086)	(0.0127)
1986			−0.0004	0.0106	0.0016	0.0070	0.0135	−0.0311
			(0.0126)	(0.0135)	(0.0087)	(0.0123)	(0.0087)	(0.0136)
1987			0.0419	0.0122	0.0213	−0.0004	0.0065	0.0313
			(0.0141)	(0.0148)	(0.0097)	(0.0138)	(0.0090)	(0.0141)
1988			−0.0056	−0.0171	0.0048	−0.0104	−0.0073	−0.0217
			(0.0155)	(0.0165)	(0.0103)	(0.0146)	(0.0093)	(0.0151)
1989			0.0852	−0.0006	0.0263	−0.0025	0.0414	−0.0036
			(0.0255)	(0.0163)	(0.0151)	(0.0153)	(0.0156)	(0.0155)
1990			−0.0383	0.0147	−0.0074	0.0224	−0.0091	0.0271
			(0.0273)	(0.0166)	(0.0154)	(0.0159)	(0.0175)	(0.0161)
1991			0.0668	−0.0220	0.0614	−0.0379	−0.0081	−0.0204
			(0.0271)	(0.0181)	(0.0224)	(0.0172)	(0.0137)	(0.0175)
1992			0.1136	−0.0080	−0.0144	0.0219	0.0650	−0.0187
			(0.0659)	(0.0172)	(0.0266)	(0.0167)	(0.0269)	(0.0178)
	$\chi^2_{17} = 81.05$		$\chi^2_{24} = 65.38$		$\chi^2_{24} = 34.46$		$\chi^2_{17} = 27.55$	
	$P = 0.000$		$P = 0.000$		$P = 0.077$		$P = 0.051$	

Estimates are calculated from (14), (15), and (16) using log expenditure and log income per equivalent adult. Associated χ^2 tests are presented below.

APPENDIX 6: MINIMUM DISTANCE ESTIMATES OF CHANGES IN THE VARIANCES OF TRANSITORY AND PERMANENT SHOCKS TO INCOME USING LEVELS

Year	Born 1920s $\Delta\,\mathrm{var}_{kt}\,(u)$	$\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1930s $\Delta\,\mathrm{var}_{kt}\,(u)$	$\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1940s $\Delta\,\mathrm{var}_{kt}\,(u)$	$\Delta\,\mathrm{var}_{kt}\,(v)$	Born 1950s $\Delta\,\mathrm{var}_{kt}\,(u)$	$\Delta\,\mathrm{var}_{kt}\,(v)$
1969	0.0018		−0.0016		0.0104			
	(0.0159)		(0.0176)		(0.0247)			
1970	0.0240	0.0063	0.0157	−0.0227	−0.0050	0.0147		
	(0.0199)	(0.0238)	(0.0201)	(0.0245)	(0.0247)	(0.0283)		
1971	0.0301	−0.0441	0.0197	−0.0015	0.0275	−0.0031		
	(0.0221)	(0.0280)	(0.0224)	(0.0293)	(0.0251)	(0.0292)		
1972	−0.0094	0.0618	0.0121	−0.0018	0.0404	0.0074		
	(0.0218)	(0.0270)	(0.0235)	(0.0288)	(0.0255)	(0.0305)		
1973	0.0215	−0.0351	−0.0264	−0.0183	−0.0339	−0.0144		
	(0.0240)	(0.0298)	(0.0233)	(0.0302)	(0.0259)	(0.0311)		
1974	−0.0215	−0.0080	0.0275	0.0112	−0.0191	−0.0308		
	(0.0248)	(0.0342)	(0.0245)	(0.0308)	(0.0255)	(0.0326)		
1975	−0.0043	−0.0316	−0.0349	0.0163	0.0025	0.0180		
	(0.0231)	(0.0372)	(0.0231)	(0.0326)	(0.0247)	(0.0327)		
1976	−0.0117	0.0321	0.0151	−0.0606	−0.0150	−0.0152	−0.0131	
	(0.0212)	(0.0329)	(0.0201)	(0.0331)	(0.0234)	(0.0321)	(0.0342)	
1977	−0.0147	−0.0338	−0.0337	0.0453	−0.0412	0.0249	−0.0254	0.0319
	(0.0191)	(0.0308)	(0.0190)	(0.0301)	(0.0203)	(0.0306)	(0.0270)	(0.0415)
1978	0.0579	0.0868	0.0546	0.0251	0.0653	0.0064	0.0138	0.0152
	(0.0214)	(0.0286)	(0.0222)	(0.0288)	(0.0230)	(0.0316)	(0.0272)	(0.0402)
1979	0.0217	−0.0203	0.0245	0.0052	−0.0145	0.0062	0.0311	0.0164
	(0.0267)	(0.0357)	(0.0270)	(0.0342)	(0.0264)	(0.0350)	(0.0301)	(0.0426)
1980	0.0643	−0.0473	0.0218	−0.0314	0.0341	−0.0094	0.0763	−0.0392
	(0.0297)	(0.0421)	(0.0282)	(0.0398)	(0.0277)	(0.0382)	(0.0339)	(0.0476)
1981	0.0364	0.0360	0.0282	0.0185	−0.0089	−0.0076	−0.0288	0.0531
	(0.0315)	(0.0422)	(0.0301)	(0.0417)	(0.0270)	(0.0400)	(0.0315)	(0.0477)
1982	−0.0210	−0.0896	−0.0154	−0.0250	0.0101	0.0094	−0.0434	−0.1106
	(0.0324)	(0.0418)	(0.0298)	(0.0416)	(0.0257)	(0.0411)	(0.0314)	(0.0477)
1983	0.0269	0.1164	0.0213	0.0853	0.0063	0.0013	−0.0241	0.0937
	(0.0330)	(0.0398)	(0.0313)	(0.0424)	(0.0294)	(0.0413)	(0.0291)	(0.0441)
1984	−0.0554	−0.0805	0.0070	−0.0951	−0.0465	−0.0043	0.0338	−0.0134
	(0.0317)	(0.0467)	(0.0351)	(0.0523)	(0.0289)	(0.0452)	(0.0297)	(0.0482)
1985	0.0626	0.0517	0.0927	0.0266	0.0717	0.0575	0.0282	0.0523
	(0.0344)	(0.0484)	(0.0407)	(0.0545)	(0.0312)	(0.0487)	(0.0335)	(0.0523)
1986			0.0466	0.0287	0.0478	−0.0092	0.1052	−0.0649
			(0.0501)	(0.0588)	(0.0410)	(0.0562)	(0.0415)	(0.0601)
1987			0.1732	0.0311	0.1436	0.0688	0.0902	0.1848
			(0.0642)	(0.0738)	(0.0547)	(0.0705)	(0.0513)	(0.0724)
1988			0.0967	−0.0270	−0.0101	−0.1371	0.0323	−0.1896
			(0.0733)	(0.0859)	(0.0580)	(0.0819)	(0.0581)	(0.0864)
1989			−0.0001	−0.0866	0.0500	0.0589	0.0672	−0.0093
			(0.0730)	(0.0848)	(0.0552)	(0.0848)	(0.0624)	(0.0877)
1990			0.0922	0.1457	0.1956	0.1370	0.0207	0.2001
			(0.0840)	(0.0906)	(0.0709)	(0.0960)	(0.0677)	(0.0935)
1991			0.0079	−0.1526	0.0475	−0.3522	0.0178	−0.1875
			(0.0897)	(0.1093)	(0.0823)	(0.1098)	(0.0726)	(0.1120)
1992			0.0272	−0.0309	0.0287	0.1975	0.0149	−0.0862
			(0.0837)	(0.1003)	(0.0794)	(0.1000)	(0.0703)	(0.1134)
	$\chi^2_{17} = 158.5$		$\chi^2_{24} = 212.9$		$\chi^2_{24} = 105.4$		$\chi^2_{17} = 31.95$	
	$P = 0.000$		$P = 0.000$		$P = 0.000$		$P = 0.015$	

Estimates are calculated from (14), (15), and (16) using levels of expenditure and income per equivalent adult. Associated χ^2 tests are presented below.

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