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Attanasio, O.P.; Weber, G.

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by Orazio P. Attanasio
and Guglielmo Weber

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ORAZIO P. ATTANASIO* and GUGLIELMO WEBER**[1]

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* Stanford University and
Centre for Labour Economics

** University College London and
Institute for Fiscal Studies

ABSTRACT

In this paper we try to asses the size of the elasticity of intertemporal substitution (EIS) and to explain why very different results are obtained using aggregate and micro data. We use two overlapping generations models to establish that estimates of the EIS based on aggregate data are downward biased due to the presence of cohort effects and productivity growth.

These predictions are confirmed by our empirical results. Our estimate of the EIS using average cohort data is just above unity and is reasonably well determined. The estimates we get using different measures of aggregate data (either from National Account statistics or average Survey data) are instead consistently lower.

Our theoretical model can also be used to explain the "excess sensitivity" of aggregate consumption to labour income, even in a world without liquidity constraints or myopia. We estimate an equation similar to that recently proposed by Campbell and Mankiw (1987) and obtain mixed results.

While our empirical estimates must be taken with some caution, we hope to have proved that the investigation of pseudo panel data based on Expenditure Survey can give considerable insights on consumption behaviour and dominates the use of aggregate National Account data.

Introduction

The relation between consumption growth and the interest rate has recently been the object of renewed interest.[1] It has been argued that regressing consumption growth on the interest rate, provided that simultaneity problems are properly allowed for, one should obtain an estimate of the elasticity of intertemporal substitution.[2] A proper evaluation of such an elasticity is, needless to say, extremely interesting and important for many theoretical and policy problems. Nonetheless there is no agreement in the profession about the actual size of such a coefficient. Mankiw, Rotemberg and Summers (1985) obtained a high estimate, while Hall (1988) claims that their result is biased by the use of inappropriate instruments and that the coefficient is actually very low.

In a previous paper (Attanasio and Weber, 1989) we obtained very different results using aggregate consumption and average cohort data: the estimate of the elasticity of intertemporal substitution based on aggregate data was much lower than that based on cohort data. In this paper we try to explain this difference, and argue that the use of aggregate national account data is inappropriate. Thee are two basic reasons for this. On the one hand, the use of aggregate consumption data involves the consideration of a sizeable

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[1] See, for instance, Attanasio and Weber (1989), Hall (1988).

[2] If one assumes that a representative agent maximizes expected utility, and consumption growth and interest rates are jointly log-normal, then the coefficient of a regression of consumption growth on the interest rate can be interpreted either as the elasticity of intertemporal substitution or as the (reciprocal of) the coefficient of relative risk aversion. Under different and more general approaches such a coefficient measures the elasticity of intertemporal substitution. See Attanasio (1988a), Attanasio and Weber (1989), Epstein and Zin (1987), Hall (1988).

proportion of the population which is likely to be liquidity constrained. On the other hand we argue that the Euler equations for each generation, because of entries and exits into and from the consumption pool, do not, in general, aggregate.

The use of cohort data eliminates the latter problem and might alleviate the former.

In the first part of the paper we present two simple overlapping generation models which try to assess the direction and the size of the bias induced by entries into and exits from the consumption pool. We argue that such a bias might be substantial.

In the second part we present some empirical evidence which confirms our previous results: the use of aggregate data (either national account or average survey data) gives us much lower estimates of the elasticity of intertemporal substitution. Particular care is devoted to the choice of appropriate instruments for our processes. As Hall (1988) pointed out, time aggregation would cause MA residuals which would make lagged one instruments invalid. Furthermore, working with rates of growth of average cohort data, would again give, because of measurement error, MA residuals (see Deaton, 1985). This second source of residual autocorrelation seems to prevail in our results.

To investigate the relative importance of our two explanations of a lower estimate of intertemporal substitution with aggregate data, we also investigate what are the effects of adding to our "base" cohort younger people (which are more likely to be subject to liquidity constraints) and what are the effects of using age -band data from the survey we used to get the cohort data. In a recent paper Campbell and Mankiw (1987) test for the excess sensitivity of consumption to income by adding income variable to the regression of consumption on the interest rate. We perform the same test for our cohort data.

The paper is organized as follows. In Section 2 we present two overlapping generation models and some simulations of their dynamic behavior. In Section 3 we discuss the data and the econometric techniques used on the paper . In Section 4 we present the empirical results. Section 5 concludes the paper. In the appendix we derive the variance covariance ratio of the IV estimator discussed in section 3 of the paper.

2. Cohort and Aggregate Consumption in an Overlapping Generation Framework.

2.1 A Two-Period Model.[3]

Consider an economy populated of overlapping generations of L identical individuals who live for two periods.[4] In the first period of their life they inelastically supply their labour endowment, and are paid the current wage. In the second period they retire and live off what they saved in the first period. The capital stock, which is used with labour by a constant return to scale production function, is the only way of transferring resources to the future. Such a capital stock is assumed to depreciate completely in the production process. This assumption is not necessary in the two period model, but it will be in the three period one.

The individuals are assumed to maximize expected utility; the instantaneous utility function is logarithmic. Therefore at time t an individual of age 1 will face the following problem:

$$(1) \quad \text{Max } \log c_t^1 + \beta E_t \log c_{t+1}^2$$

subject to:

$$(2) \quad c_t^1 + s_t^1 = w_t$$

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[3] The model presented here is basically Diamond (1965) model without government bonds and with a particular type of uncertainty. The result we derive is similar to the one obtained by Blanchard (1985) in a model where the only source of uncertainty is the timing of death.

[4] The two period model and the three period model below could be easily generalized to allow for population growth. We do not do it here for notational simplicity.

and

(3) $s_t^1\ (1+r_{t+1}) = c_{t+1}^2$

where c_t^i is consumption of individuals of age i at time t, s_t^1 is the saving of the young at time t, and w_t and r_t are the wage and the interest rate respectively.

The only uncertainty in this economy comes from production. Output, which under perfect competition is distributed to the factors of production according to their marginal productivity, is given by a CES production function with a multiplicative, Hicks- neutral, productivity shock, whose log is a random walk. Technical progress is deterministic and Harrod- neutral.[5]

Let Y_t be output at time t, L_t the number of workers and K_t the capital stock. Aggregate production is :

(4) $Y_t = Z_t\ [\alpha\ K_t^{-\rho} + (1-\alpha)(g_t L_t)^{-\rho}]^{-1/\rho}$

where g_t is labour augmenting technical progress. If we express equation (4) in efficiency unity terms we have:

(5) $y_t = Z_t[\alpha\ k_t^{-\rho} + (1-\alpha)]^{-1/\rho}$

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[5] This specification might seem a bit clumsy. Harrod neutrality of technical progress is assumed to guarantee the existence of a well defined steady state. On the other hand the stochastic component of productivity has to be Hicks neutral to ensure the existence of a closed form solution for consumption. With such a specification, however, total factor productivity (as measured for example by Solow residuals) follows a random walk with drift.

where $y_t = Y_t/L_t g_t$ and $k_t = K_t/L_t g_t$.

It is well known that with these restrictive assumptions on tastes and technology it is possible to derive the closed form solution for consumption. It is also well known that with a Cobb-Douglas production function the individual Euler equations aggregate perfectly. In this case the use of aggregate consumption data would not bias inferences on the elasticity of intertemporal substitution.

In general the Euler equation holds for half the population: the cohort individuals who are aged 2 at t+1 and were therefore aged 1 at t. With a Cobb Douglas production function technical progress is fully neutral and will be equally shared among factors of production. This implies that the ratio of consumption of the new entrants into the consumption pool to that of the individuals who died at t behaves as consumption growth over the life cycle of the generation alive at time t and t+1: with obvious notation, ${c^1}_{t+1}/{c^2}_t = {c^2}_{t+1}/{c^1}_t$. The Euler equation will therefore hold for aggregate consumption. However, as soon as we relax the assumption of unity elasticity between capital and labour in the production function, this will no longer be true.

The closed form for consumption that can be derived from equation (1) to (3) is the following:

$$(6) \qquad c_t^1 = \frac{1}{1+\beta} w_t$$

$$(7) \qquad c_t^2 = \frac{\beta}{1+\beta} w_{t-1}(1+r_t)$$

where w_t is the wage rate and r_t is the interest rate. Note that from (6) and (7) it follows that the Euler equation $c^2_{t+1}/c^1_t = \beta(1+r_{t+1})$ will hold exactly.

From the first order condition for profit maximization we know that:

$$(8) \qquad w_t/g_t = \frac{\partial Y_t}{\partial L_t} = (1-\alpha)Z_t^{-\rho}\, y_t^{1+\rho}$$

$$(9) \qquad 1+r_t = \alpha\, z_t^{-\rho}\, y_t^{1+\rho}/k_t^{1+\rho}$$

From this it follows that:

$$(10) \qquad \frac{c^1_{t+1}}{c^2_t} = \frac{1}{1+\beta}\;\frac{(1-\alpha)}{\alpha}\;\left(\frac{y_{t+1}}{y_t}\right)^{1+\rho}\left(\frac{Z_{t+1}}{Z_t}\right)^{-\rho} k_t^{\rho}$$

$$= (1+r_{t+1})(k_{t+1}k_t)^{\rho}\left(\frac{1-\alpha}{\alpha}\right)^2 \frac{(1+\beta)^2}{\beta}$$

If we define aggregate consumption as the geometric mean of c^1_t and c^2_t we have that

$$(11) \qquad \Delta \log c_{t+1} = \text{constant} + \log (1+r_{t+1}) + \frac{1}{2}\rho\, [\log (k_{t+1}) + \log (k_t)]$$

For $\rho = 0$, in which case the CES reduces to the Cobb-Douglas, equation (11) shows that the Euler equation holds in the aggregate. However if $\rho > 0$ (capital and labour are worse substitute than in the Cobb Douglas case) there will be a downward bias in the estimation of the elasticity of intertemporal substitution, if aggregate data is used.

Equation (11) is a deterministic equation which holds with no error. The reason for this lies in the particular assumption on uncertainty which was made to obtain a closed form solution for consumption. However it can still be interpreted as a regression equation with the variance of the residuals equal to zero.

In general, to evaluate the bias deriving from the use of aggregate data, one should project aggregate consumption growth on the interest rate. In this case, with a zero variance residual, this is equivalent to evaluating the derivative of consumption growth with respect to the interest rate[6]. This can be done using equations (8) and (9).

$$\frac{\partial \Delta \log c_{t+1}}{\partial \log (1+r_{t+1})} = 1 + \frac{1}{2}\rho\left\{\left[(1+\rho)'\left(\frac{\partial \log y_{t+1}}{\partial \log k_{t+1}} - 1\right)\right]^{-1} - \left[(1+\rho)^2 \frac{\partial \log y_t}{\partial \log k_t}\left(1 - \frac{\partial \log y_{t+1}}{\partial \log k_{t+1}}\right)\right]^{-1}\right\} \tag{12}$$

[6] Another way of seeing the same problem is as an omitted variable problem. In a more general setting one would look at population correlations, in this case, the analysis of the derivative is legitimate.

$\partial \log y_{t+1}/\partial \log k_{t+1}$, is the share of capital in value added and is therefore less than unity. Hence the bias is unambiguously negative.

The elasticity of aggregate consumption to the interest rate will be further away from unity the higher is ρ.

This shows that the bias introduced by considering aggregate consumption data to estimate the elasticity of intertemporal substitution from an Euler equation, can be substantial. In the economy we have been considering, entries and exits into and from the consumption pool account for half the population. Furthermore, the new entrants have access to higher lifetime resources because of productivity growth. Aggregation fails even in this two period model because technical progress is not neutral when the production function is not a Cobb Douglas.

Already implicit in this argument there is a possible objection, in defense of aggregate data: if one uses, as we do, quarterly data, one should consider an overlapping generation model with 160-200 periods. Therefore the entrants will account for a much smaller proportion of the population. There are several answers to this argument. We have already noticed that the bias from using aggregate data will be higher, the higher is ρ, i.e., the less substitutability there is between labour and capital. As a first approximation it seems safe to say that, on a quarterly basis, capital and labour are not very good substitutes and surely less so than on a 10 years basis. It has also to be said that with more than one period even with log utility, the effects of the interest rate on the level of consumption are much more complex than in the two period model, so that it is not clear that the bias in equation (11) will simply be proportional to $\rho/160$ instead of $\rho/2$. Finally, the most convincing argument about the importance of entries and exits for the aggregate relationship between consumption growth and the interest rate is the one made by Modigliani and Brumberg (1954) and stressed by Deaton (1987). Consider an overlapping generations economy and assume that each generation lives a large number of periods. Consider a stationary steady state for such an economy, so that aggregate consumption does not grow. Suppose also that , in the steady state, the interest rate is greater than the discount rate. This will imply that individual consumption will be growing over the life cycle. If we consider two stationary steady states with two different interest rates, we will observe a different rate of growth of individual consumption , and still , at the aggregate level, we will not be able to detect any relation between consumption growth and the interest rate. The model above and the one that follows can be interpreted as formalizations and generalizations of the this intuition, and as ways to evaluate the bias induced in the estimation of the elasticity of intertemporal substitution by the consideration of aggregate data as opposite to cohort data.

To investigate the effects of richer dynamic structures we now analyze a three period model.

2.2 Three Period Model

We now consider overlapping generations of L identical individuals who live 3 periods: they work in the first two and retire in the third. All the assumptions about tastes, technology, productivity growth are the same as in the two period model. Nonetheless the solution of the model becomes much more problematic. It will be necessary to assume full depreciation of the capital stock to be able to obtain a sort of closed form solution for consumption.

The consumer problem is solved backward starting with the last period.

In the third period of his life the choice is trivial: he will consume what his savings yield. In the second period of his life, given what he saved in the first period, he has to decide how much to consume. The problem is very similar to that of the two period model, with the difference that he will receive not only labour income, but also capital income. The solution for consumption is however straightforward and it will be given by:

$$(13) \quad c^2_{t+1} = \frac{1}{1+\beta} ((w_t - c^1_t)(1+r_{t+1}) + w_{t+1})$$

Given equation (13), if we go back to period one we will have the following maximization problem as of time t:

$$(14) \quad \text{Max } E_t \log (c^1_t) + \beta \log (c^2_{t+1}) + \beta^2 \log (c^3_{t+1})$$

subject to equation (13) and to:

$$(15) \quad c^3_{t+2} = \frac{\beta}{1+\beta} [(w_t - c^1_t)(1+r_{t+1}) + w_{t+1}](1+r_{t+2})$$

Substituting (13) and (15) into (14) we get:

$$\text{Max } \log (c^1_t) + \beta E_t \{ \log [(w_t - c^1_t)(1+r_{t+1}) + w_{t+1}] +$$

$$+ \beta \log [(w_t + c^1_t)(1+r_{t+1}) + w_{t+1}] + \beta \log (1+r_{t+1}) +$$

$$+ \beta \log (\frac{\beta}{1+\beta}) + \log (1+\beta) \}$$

Taking the derivative of (16) with respect to c_{1t} yields:

$$(17)\quad \frac{1}{c_t^1} = E_t \left\{ (\beta+\beta^2) \frac{1}{w_t - c_t^1 + \frac{w_{t+1}}{1+r_{t+1}}} \right\}$$

w_{t+1} and r_{t+1} are two random variable whose distribution is, in general, unknown. However, the ratio $w_{t+1}/(1+r_{t+1})$ is <u>not</u> a random variable. From the first order condition for profit maximization it can be proved that[7]

$$(18)\quad \frac{w_{t+1}}{1+r_{t+1}} = \frac{1-\alpha}{\alpha} k_{t+1}^{-\rho}$$

k_{t+1} is a variable known at time t because is completely determined by the saving decisions taken at time t. We can therefore eliminate the expectation operator from equation (17) and write the following expression for consumption in the first period.

$$(19)\quad c_t^1 = \frac{1}{1+\beta+\beta^2} \left(w_t + \frac{w_{t+1}}{1+r_{t+1}} \right)$$

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[7] It is here that the assumption of full depreciation of capital plays an important role. Without it the ratio $w_{t+1}/(1+r_{t+1})$ would still be a random variable and it would be impossible to eliminate the expectation and obtain a closed form solution for consumption. An implicit assumption used here as well as in the previous model, is that individuals are price takers in the capital market: they consider their saving decisions as having negligible effect on the interest rate (and the wage rate).

Note that, even though we still have a logarithmic utility function, consumption (unlike in equation (6)) does depend on future interest rate because people work for two periods. Equation (19) is not strictly speaking a closed form solution because k_{t+1} will depend (in a non-linear fashion) upon c_t^1 (and on c_t^2). However such an expression is sufficient to study the bias implied by the use of aggregate data.

It is easily checked that equations (13), (15) and (19) satisfy the following Euler equation.

$$(20)\quad E_t[(c^i_{t+i-1}/c^1_t)^{-1} \prod_{n=0}^{i-2} (1+r_{t+1+n})\beta^{i-1}] = 1 \quad , \quad i = 2,3$$

The model is then completed by the following identities:

$$(21)\quad \frac{1}{2}(c_t^1 + c_t^2 + c_t^3) = w_t + r_t k_t - (k_{t+1} - k_t)$$

$$(22)\quad \frac{1}{2}(k_t^2 + k_t^3) = k_t$$

Equation (21) is the national account identity. k^2 and k^3 represent the capital stock held by an individual of age 2 and 3 respectively. The left hand side is multiplied by 1/2 because consumption is expressed in per capita terms, while k is expressed in per worker terms and there are $2L$ workers and only L individuals in each cohort. Capital accumulation is governed by the following equations:

$$(23)\quad k_t^2 = w_{t-1} - c_{t-1}^1$$

(24) $k^3_t = w_{t-1} - c^2_{t-1} + (1+r_{t-1})k^2_{t-1}$

It is easy to show that, once again, the Euler equation does not hold in the aggregate because c^1_{t+1}/c^3_t is not equal to c^3_{t+1}/c^2_t or to c^2_{t+1}/c^1_t. However the bias is this time much more difficult to evaluate because of the extra interest rate effects present in consumption.

To assess the magnitude of the bias and compare it to the one in the two period model, we simulated the two models with some plausible values for the parameters in the production function and the discount rate β[8]. We started the simulations from the steady state and generated a vector of 100 random shocks Z such that $\log(Z_t) = \log(Z_{t-1}) + e_t$, with e distributed as a normal variable with variance σ^2. g_t is assumed to grow at a rate μ. It can be easily shown that our production function implies that total factor productivity evolves as a random walk with drift, where the drift term is μ .

In the two period model we could solve for all the relevant variables analytically, while in the three period model we had to use a numerical method to solve equation (19) to get an expression for consumption of people aged 1 in each period.

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[8] A surprising feature of these OG models with a production function less elastic than the Cobb Douglas is that for many plausible values of the parameters, no steady state (other than the trivial one with everything equal to zero) exists. In general the lower the elasticity of substitution between capital and labour and the higher α (the parameter which governs the share of capital) the less likely is the steady state to exist. In the two period case, for an elasticity of substitution between capital and labour of 0.5, values of α greater than .1 prevent the existence of the steady state. In the three period model the situation is better which seems to indicate that the more periods one considers the more likely is the steady state equilibrium to exist. See Attanasio (1988b).

Having generated all the data, we 'estimated' the elasticity of intertemporal substitution in the two models using aggregate consumption data projecting aggregate consumption growth on the interest rate. We repeated this procedure 150 times.

The results of this exercise are reported in table 1. The first panel reports the average estimate of the elasticity of intertemporal substitution obtained defining aggregate per capita consumption as the arithmetic mean of consumption across cohorts, while in the second panel it is defined as the geometric mean. For the two period model we report the results obtained with an average productivity growth of 20%. The bias is an increasing function of average productivity growth: larger productivity growth implies a larger difference between the life time resources of new entrants and the people who die in the previous period. For the three period model we report the results with average productivity growth of 13.3% . The standard deviation of the productivity shock is such that its coefficient of variation on an annual basis is equal to one[9].

The justification for the lower productivity growth in the three-period is the fact that going from two to three periods we narrow the length of each period: 20% is the growth per period which would correspond to 13.3% if the period was 3/2 longer. These figures are quite conservative : if we think of a life cycle of 50 years, a rate of productivity growth of 2% per year would imply a drift of 50% in the two period model and of 33% in the three period one .

[9] The standard deviation for the random variable z has been calibrated looking at the standard error of the Solow residuals for value added in the UK for the period 1952-1984.

The first thing to note is the size of the bias in both models. As expected the bias in the three period model is smaller than in the two period one. However it is larger than two thirds of the bias in the two period model. Furthermore, we did not decrease the elasticity of substitution between labour and capital going from two to three periods. This will reinforce the size of the bias.

From this experiment it can be concluded that aggregation can represent an important source of bias in the estimation of the elasticity of intertemporal substitution. In the next section we investigate this proposition empirically estimating this parameter both on cohort data and on data aggregated over different cohorts.

3. Data and estimation issues.

In our empirical application we have used British data drawn from the National Accounts (CSO Data Bank) and from the Family Expenditure Survey. While the first source is commonly employed in applied macro-economic work the latter has been employed mainly by micro-researchers, thus deserving some illustration.

The British Family Expenditure Survey is run every year on a randomly selected sample of around 7,000 households. The original sampling design covers about 10,000 households, hence little less than a third of the households fails to respond. Non-response includes cases where the interviewer failed to make contact, cases where at least one adult member refused to answer a question or to fill in his/her diary and cases where the quality of the information given was very poor (inconsistent answers). It should be stressed that participation in the survey is voluntary and no replacement takes place if the interviewer fails to make contact.

The British FES has been extensively studied and used by economists in the past. Atkinson and Micklewright (1983), e.g., thoroughly addressed the issue of grossing up income data from the FES, and reached the conclusion that the only types of income where under-reporting is substantial are investment and self-employment income. On consumption data, under-reporting is noticeable only on alcohol, a relatively small item. Expenditure on other items is thought to be accurately recorded, thanks to the careful sampling design. Each household is in fact interviewed twice and is asked to produce the latest gas, electricity and telephone bills, records of rent paid, of lumpy purchases and of any credit-financed expenditure. On top of answering questions on work-related matters and on expenditure on durable goods, each adult household member fills in a detailed diary over a two-weeks period.

In our application we use fifteen years (1970-84) of FES data to construct quarterly series on consumption, income and prices. For each household we take total expenditure on non durable goods and services (exclusive of housing services), net household income (inclusive of earnings, self-employment income, pensions, social security ; exclusive of investment income and imputed rent from owner occupation) and a household specific Stone price index based on the same eight broad groups of commodities included in the consumption measure and on their published retail price indices . We then take geometric means of each variable over each quarter and construct the following pseudo-aggregate data:

1) average cohort data, where the cohort includes all those households whose head was born in the 1930-40 interval;

2,3) two average age-band data[10], where the age bands include all households whose head's age is 35-50 in the first case and 30-55 in the second case;

4) average FES data, including all participant households.

For consistency with aggregate National Accounting data we also produce average FES data by taking simple arithmetic means. Both consumption and income variables are expressed in per-capita terms, where the denominator is the total number of household members, i.e. it includes both adults and children.

The aggregate data we use is as close as possible to the survey data available. In particular, we take consumers' expenditure over the same set of commodities and use personal disposable income as consumption and income

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[10] The average age band data are constructed taking, for each (time) observation , every household in the cohort whose age is included in the age band. Entries and exits will therefore be proportionally more important when the age band is narrower.

variables, respectively. The price index we adopt is the appropriate implicit deflator; we also experimented with an RPI-based index and found no substantial differences.

As for asset returns data, we have chosen to operate with Building Societies Deposits. This type of interest-bearing deposit is particularly attractive on a number of grounds. First of all, it was the most commonly held asset over the sample period (over 50% of FES respondents quoted interest income of this type). Secondly, the quoted interest rate is net of tax for standard rate tax-payers (the great majority of households). Thirdly, it is an asset where negative holdings are common, in the form of mortgages: tax-exemption rules imply that for many households after-tax lending and borrowing rates are very close to each other (in 1984 the interest paid on the first 30,000 pounds of most mortgages was tax-deductible. Typically, only young households would have a larger outstanding mortgage.).

The equation we estimated for the above mentioned data sets is the following:

(25) $\log(c_{t+1}/c_t) = \text{constant} + \sigma\log(1+r_{t+1}) + e_{t+1}$

This equation can be derived, assuming joint conditional lognormality of consumption growth and the real interest rate, from the Euler equation for consumption. The parameter σ should be interpreted, as Hall (1988) convincingly argues , as the elasticity of intertemporal substitution[11] .

An econometric issue is worth discussing. No matter how good the quality of the data used, individual households data are likely to be affected by measurement error. Aggregate data are the result of averaging over a very large

[11] Under expected utility and time separability σ is also equal to the reciprocal of the coefficient of relative risk aversion.

number of households and measurement error is likely to disappear. In average survey data measurement error is likely to persist, unless the number of included households is very large[12]. However, some information on its structure and size is available. In equation (25) the disturbance e_{t+1} will be the sum of the expectational error (which is either white noise or an MA(1) with positive coefficient, Hall, 1988) and of the measurement errors of consumption growth and the inflation rate[13], which are MA(1) processes with negative unitary coefficient. The sum of two MA(1) processes (or of an MA(1) and a white noise) is still an MA(1). Because the two original MA(1) -the one originated by the measurement error and the one originated by time aggregation- have coefficients of opposite sign, the first order autocorrelation of the resulting MA can be of either sign, depending on the relative variances of the original processes. Our data seems to indicate negative first order autocorrelation, thus suggesting that measurement error is more important than time aggregation in this data set.

Given the presence of MA residuals the use of instruments dated t-1 is invalid, but instruments lagged 2 and more will yield a consistent estimator. However, given the linearity of equation (25), the sample information on within cohort variances and covariances of the various variables, can be used to implement a more efficient estimation technique, as suggested by Deaton (1985).

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[12] The implicit assumption here , and in the discussion that follows, is that measurement error (or the idiosyncratic component of household consumption) is independently distributed across households. This is an identifying restriction which cannot be tested in the present framework.

[13] Besides consumption growth the real interest rate (and all the instruments used below) is also affected by measurement error. This is because the price index used to compute the inflation rate is an average of household specific price indexes. Measurement error seems to be particularly important for reported labour income.

One can show that in the case in hand the IV estimator requires the selection of instruments lagged 2 or more and involves simple reweighting of the projection matrix so as to give more weight to the instruments less affected by measurement error. Given the amount of noise present in the data at the household level, taking into account measurement error explicitely can imply a considerable efficiency gain. The computation of the variance - covariance matrix of such an estimator requires a further correction to account for the serial correlation in the disturbance[14] . The derivations are presented in the Appendix[15].

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[14] The presence of MA residuals does not affect the consistency of the standard IV estimator (provided instruments dated t-2 are chosen), but the estimates of the standard errors will be biased.

[15] The corrections are made under the assumption that the variance-covariance matrix of the measurement errors is known. This hypothesis is relatively harmless given that such a matrix is estimated using about 17,000 observations.

4. Estimation results.

We have estimated an Euler equation for consumption for each of the data sets discussed in section 3. The results are reported in table 2. We present both estimates obtained using standard IV and those that correct for the presence of measurement error. As already said, the IV estimates presented in this table are derived using an instrument set including variables lagged 2 or more. The instrument set included lagged values of consumption growth, the expost real interest rate, growth in labour income and the inflation rate. The estimated Euler equation allows explicitly for seasonality, along the lines proposed by Miron (1986). The full set of estimates is available from the authors upon request.

We start from aggregate data, where we can confirm our previous findings of a rather low elasticity of intertemporal substitution (EIS). The estimate is well determined (0.34 ,t-value=3.6)[16]. Of course, no information is available on the measurement error for aggregate data.

In the second row of table 2 we report the estimates of the elasticity of intertemporal substitution obtained using the data for the cohort of household whose head was born between 1930 and 1940. We chose this particular cohort so that we observe the behaviour of people whose age is between 30 and 55. If we assume away migration (hardly a major feature of British society), entries and

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[16] This contrasts with our previous results of a poorly determined estimate for the aggregate data. The difference is accounted by the use of some extra instruments : the rate of growth in labour income and the inflation rate.

exits should not affect average cohort consumption. Not only, this group of households is on a stretch of their life-cycle where liquidity constraints are least likely to be present (as suggested by simulation studies, Davies, 1981, e.g., and by more direct empirical evidence, Weber, 1988; see also Alessie, Melenberg and Weber, 1988) and where the probability of death is fairly constant. Hence the EIS estimate derived from this data is our best guess of the true parameter (this statement remains correct if one believes that the dynastic view is correct).

The estimated EIS is considerably higher than the one obtained with aggregate data : the measurement error corrected point estimate is 1.30, with a t-value of 2.5. This result is consistent with the theoretical predictions of section 2.[17] The Sargan test of overidentifying restrictions does not reject the null hypothesis.

In the third row we analyze the possibility that our results are not a consequence of using cohort as opposite to aggregate data , but a peculiarity of the FES data set. This would be surprising, given that the CSO uses the FES data (along with other information) to compute estimates of the National Account data. The estimate of the EIS obtained using data from the whole FES is remarkably close to that from the aggregate data (0.42, t-value 1.6 for standard IV and 0.35, t-value 1.73 for measurement error corrected IV). We also experimented on the likely effects of taking geometric, rather than arithmetic, averages: the estimates on the whole of FES using the two different procedures were almost identical (EIS=0.40, t-value 1.2).

In the fourth row we report the estimates obtained using consumption data

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[17] It is also remarkably similar to the estimates obtained by Blundell, Browning and Meghir (1989) who adopt a micro-oriented approach.

for people aged between 30 and 55 in each quarter. Entries and exits are proportionally more important in this data set, so that if our theoretical arguments in section 2 are valid we should get an estimate lower than for the whole FES[18]. This is indeed the case for the measurement error corrected IV: the EIS is estimated at 0.29 (t-value 0.97).

As a further test of the importance of the weight of entries and exits for the bias in the estimate of the EIS we repeated the exercise just discussed for the age-band 35-50. In this case the proportion of entries and exits is larger than in the 30-55 case , so that we would expect a lower estimate. This is not the case for the estimates reported in table 2. However, given the imprecision of the estimates, this last experiment is not conclusive: the standard errors are sufficiently large to make point estimates uninformative.

In a recent paper Campbell and Mankiw (1987) have proposed a test of the hypothesis of liquidity constraints. They argue that if a constant proportion q of the population is liquidity constrained and consumes its labour income, while for the others the Euler equation holds, an IV regression of consumption growth on the interest rate and the growth in labour income should reveal, through the coefficient on this last variable , the proportion of households that are liquidity constrained. In section 2 we saw that aggregation can bring about an artificial correlation between income and consumption growth even in the absence of liquidity constraints: income growth could proxy for the growth in life-time resources now available to the new entrants in the consumption pool. To test these hypotheses in table 3 we report the results of an IV

[18] This argument does not consider the possibility that the whole FES data contains consumption of people who are more likely to be liquidity constrained.

regression of consumption growth on the interest rate and income growth. The instruments used are the same as in table 2. The results are mixed and of no easy interpretation.

For average cohort data the standard IV technique yield results which differ from those obtained by Campbell and Mankiw (1987) : the coefficient on the interest rate remains close to unity (its point estimate is .82) , the coefficient on labour income is 0.5 (with a t-ratio of 2.4) . However , time aggregation and small cohort size imply a serially correlated error term . Under these conditions the IV estimator is inefficient and the reported standard errors are incorrectly computed. When we correct our estimates for measurement error and compute asymptotically unbiased standard errors, our results look rather different. In particular the coefficient on labour income is large (1.02) but statistically insignificant (t-value 1.4).

If we use the whole FES, instead, we find clear evidence of excess sensitivity of consumption to income : the coefficient on the interest rate falls toward zero, that on labour income growth is large and significant. The same is true for the age-band data (particularly for the smaller band), indicating that excess sensitivity generated by entries and exits can be substantial.

For aggregate data, the coefficient on the interest rate is unaffected, while the coefficient on labour income growth is 0.24 with a t-value of 1.81. This estimate is lower than one might expect, especially given our findings on whole FES data. Part of the difference may be accounted for by discrepancies between the income definitions in the two data sources, while part could be due to special features of the FES sample (e.g. : elderly household are under-represented).

Conclusions.

In this paper we tried to asses the size of the elasticity of intertemporal substitution and to explain why very different results are obtained using aggregate and micro data. In section 2 we proposed two overlapping generations models which predict a downward bias in the estimation of the EIS using aggregate data due to the presence of cohort effects and productivity growth.

These predictions have been confirmed by our empirical results, reported in section 4. Our estimate of the EIS using average cohort data is just above unity and is reasonably well determined. The estimates we get using different measures of aggregate data (either from National Account statistics or average Survey data) are instead consistently lower.

Our theoretical model can also be used to explain the "excess sensitivity" of aggregate consumption to labour income , even in a world without liquidity constraints or myopia. We estimated an equation similar to that recently proposed by Campbell and Mankiw (1987) and obtained mixed results.

While our empirical estimates must be taken with some caution, we hope to have proved that the investigation of pseudo panel data based on Expenditure Survey can give considerable insights on consumption behaviour and dominates the use of aggregate National Account data.

Table 1

Bias in the estimate of the elasticity of intertemporal substitution in two period and three period OG models.

	Two Period	Three Period
	$\mu = 0.2$ $\sigma_z = 0.063$	$\mu = 0.133$ $\sigma_z = 0.052$
geom.	0.0377 (0.9623)	0.2431 (0.7569)
aritm.	0.0417 (0.9583)	0.2421 (0.7579)

$\alpha = 0.1$
$\beta = 1.0$
$p = 1.0$

Table 2.

Estimates of the elasticity of intertemporal substitution (eis) from different data sets.

	eis (t-value)	S^b (dof)	eis[a] (t-value)	$S^{a,b}$ (dof)
Aggregate data[c]	0.3389 (0.0952)	12.40 (11)	- -	- -
Average Cohort Data : Aged 30-40 in 1970	1.3874 (0.5577)	6.19 (11)	1.3002 (0.5147)	5.59 (11)
Whole FES data	0.4234 (0.2664)	10.3 (11)	0.3514 (0.2026)	10.3 (11)
Age Band[d] data: 30-55	0.5281 (0.8036)	11.3 (11)	0.2867 (0.2958)	8.78 (11)
Age Band[d] data: 35-50	0.6547 (0.4367)	8.94 (11)	0.5722 0.3021	11.40 (11)

Notes to table 2: a) Estimates corrected for measurement error.

b) S is the Sargan test for overidentifying restrictions.

c) Aggregate data are taken from the National Account Statistics.

d) The 'Age Band' data is constructed taking, for every quarter, average per capita consumption of all the household whose head has an age included in the age band specified.

Table 3

Test of Excess Sensitivity to Current Labour Income

	(1)	(2)	(3)	(4)	(5)	(6)
Aggregate Data	0.316 (0.15)	0.245 (0.14)	9.45	--	--	--
Average Cohort Data: Aged 30-40 in 1970	0.822 (0.46)	0.500 (0.20)	6.1	0.540 (0.54)	1.016 (0.73)	4.0
Whole FES Data	0.170 (0.24)	0.735 (0.30)	8.8	0.011 (0.22)	0.845 (0.31)	7.8
Age Band Data: 30-55	0.069 (0.37)	0.785 (0.29)	8.6	0.028 (0.57)	0.498 (0.42)	10.6
Age Band Data: 35-50	0.268 (0.39)	0.491 (.21)	8.2	0.200 (0.29)	0.555 (0.27)	9.9

Notes to Table 3:
Columns (1), (2) and (3) contain standard IV estimates, while (4), (5) and (6) contain measurement-error-corrected estimates. Column (1) and (4) report the estimate for the coefficient on the interest rate, (2) and (5) the coefficient on labour income growth, and (3) and (6) the Sargan test of overidentifying restrictions (which is distributed as a chi-squared with 10 degrees of freedom. The same notes as in Table 2 apply.

Appendix

The IV estimator used in section 3 and its variance-covariance matrix.

The motivation of the estimator used in section 3 and the derivation of its variance covariance matrix follows closely Deaton (1985).

The algebra is complicated by the fact that we have at the same time a model in first differences and the use of instrumental variables.

Let us assume that the measurement error model for the level of the variables is the following:

$$\begin{pmatrix} \bar{y}_t \\ \bar{x}_t \end{pmatrix} \sim N \begin{pmatrix} y_t^* ; & \sigma_{00} & \sigma' \\ x_t^* & \sigma & \Sigma \end{pmatrix} . \tag{1}$$

Our theoretical model will be given by the following equation:

$$\Delta y_t^* = \Delta x_t^* \beta + \epsilon_t \tag{2}$$

Here y_t^* is the (unobservable) cohort mean of (log) consumption.

Δx_t^* is the real rate of interest. This variable will be affected by measurement error because the price index used to compute the inflation rate is household specific and is then averaged over the households in the cohort.

Given (1), the observable $\Delta\bar{y}_t$, $\Delta\bar{x}_t$ variable will be linked to the unobservable "starred" variables by the following model

$$\begin{pmatrix} \Delta\bar{y}_t \\ \Delta\bar{x}_t \end{pmatrix} = \begin{pmatrix} \Delta y_t^* \\ \Delta x_t^* \end{pmatrix} + \begin{pmatrix} v^1{}_t \\ v^2{}_t \end{pmatrix} \tag{3}$$

where $v_t^1 = u_t^1 - u_{t-1}^1$, $v_t^2 = u_t^2 - u_{t-1}^2$

and $u_t^1 = \bar{y}_t - \bar{y}_t^*$, $u_t^2 = \bar{x}_t - \bar{x}_t^*$

In general we know that Δx_t^* will be correlated with ϵ_t so that an instrumental variable technique is in order. ϵ_t is not necessarily a white noise process: problems like time aggregation could make it into an MA(1).

Equations (2) and (3) with the addition of an equation that gives the relation between Δx_t^* and the instruments to be used, can be described by the following statistical model.

$$E(m_{\Delta x \Delta y}) = \Omega\beta + 2\beta\sigma + \sigma_s \qquad (4)$$

$$E(M_{\Delta x \Delta w}) = \pi\Omega \qquad (5)$$

$$E(m_{\Delta w \Delta y}) = \Omega_1 \pi' \beta \qquad (6)$$

$$E(M_{\Delta w \Delta w}) = \Omega_1 + 2\Sigma_w \qquad (7)$$

$$E(m_{\Delta y \Delta y}) = \beta' \Omega_1 \beta + 2\sigma_{00} + \sigma_\epsilon^2 + 2\beta'\sigma_s \qquad (8)$$

Here σ_s is the covariance between Δx^* and ϵ. Ω is the matrix of second moments of Δx^*, w_t is vector of instruments, affected by measurement error, whose unobservable component ${w^*}_t$ is related to the variables to be instrumented through relation (5). Σ_w is the variance covariance matrix for the measurement error in the vector of instruments w_t. The measurement error in the instruments is assumed to be uncorrelated with the measurement error in the x's and y's. This assumption, which enormously simplifies the derivation below, is completely harmless in our application, given that we use instruments dated t-2.

Following standard arguments for the derivation of the IV estimator we have that

(9) $\hat{\beta}^{IV} = [M_{\Delta x\Delta w}(M_{\Delta w\Delta w}-2\Sigma w)^{-1}M'_{\Delta x\Delta w}]^{-1}M_{\Delta x\Delta w}(M_{\Delta w\Delta w}-2\Sigma w)^{-1}M_{\Delta w\Delta y}$

hence

(10) $\hat{\beta}^{IV}-\beta = (\Sigma_{\Delta x\Delta w}\Omega_1^{-1}\Sigma_{\Delta x\Delta w})^{-1}[M_{\Delta x\Delta w}(M_{\Delta w\Delta w}-2\Sigma_w)^{-1}(m_{\Delta w\Delta y}-M_{\Delta x\Delta w}\beta)]$

$$+ R + 0_p(T^{-1})$$

where R is a non-stochastic term.

We therefore need to determine

$$TV(m_{\Delta w\Delta y}-M_{\Delta x\Delta w}\beta) = TV(\Delta w'u)$$

where u is the equation error which involves both ϵ and the measurement error. u will be an MA(1).

Let us denote the first differences in the instruments as G; $G = \Delta w$.

Define $$\theta_i = \frac{1}{T}\sum_{t=1}^{T}(g_{it})u_t = \frac{1}{T}\sum_{t=1}^{T}(g_{it}^* + \nu_{it})u_t$$

As already said ν_{it} is uncorrelated with the measurement error in the y's and x's. It is also uncorrelated with ϵ_t.

$$E(\theta_i\theta_j) = \frac{1}{T^2} E\,[\sum_t \sum_s (g_{it}^* + \nu_{it})(g_{is}^* + \nu_{is})u_t u_s]$$

$$= \frac{1}{T^2} E\,[\sum_t \sum_s g_{it}^* g_{js}^* u_t u_s + \nu_{it} u_t g_{is}^* u_s + \nu_{it}\nu_{js} u_t u_s + g_{ut}^* \nu_{js} u_t u_s]$$

The g^*'s are assumed to be non-stochastic; hence the second and fourth term in the formula above are third moments of normally distributed, zero mean variables. They are therefore equal to zero.

Both the u's and the ν's are first order MA's. Hence:

$$E(\theta_i\theta_j) = \frac{1}{T} [E(g^*_{it}\ g^*_{jt}\ u_t^2) + E(g^*_{it}\ g^*_{jt-1}\ u_t u_{t+1})$$

$$+ \quad E(g^*_{jt}\ g^*_{i,t-1}\ u_t\ u_{t+1}) + E(\nu_{it}\ \nu_{jt} u_t^2) + 2E(\nu_{it+1}\ \nu_{jt}\ u_{t+1}\ u_t)]$$

Now, let σ^2_u denote the variance of u, and let σ_{u1} be the covariance between u_t and u_{t-1}.

Then

$$(11)\ E(g^*_{it}\ g^*_{jt}\ u_t^2) = g^*_i\ g^*_j\ \sigma^2_u$$

$$(12)\ E(g^*_{it}\ g^*_{jt+1}\ u_t\ u_{t+1}) = w^+_{ij}\ \sigma_{u1}$$

$$(13)\ E(g^*_{jt}\ g^*_{it+1}\ u_t u_{t+1}) = w_{ij}\ \sigma_{u1}$$

The last two terms are fourth-moments of a 4-variable normal:

$$\begin{matrix} u_t \\ u_{t+1} \\ v_{1t} \\ v_{jt} \end{matrix} \sim N \quad 0; \quad \begin{matrix} \sigma^2_u & \sigma_{u1} & 0 & 0 \\ \sigma_{u1} & \sigma^2_u & 0 & 0 \\ 0 & 0 & \sigma^2_{wi} & \sigma_{wij} \\ 0 & 0 & \sigma_{wij} & \sigma^2_{wj} \end{matrix}$$

Now $E(x_1x_2x_3x_4) = \sigma_{23}\ \sigma_{14} + \sigma_{31}\sigma_{24} + \sigma_{12}\sigma_{34}$

hence

$$(14)\ E(v_{it}\ v_{jt}\ u_t^2) = \sigma_{wij} \cdot \sigma^2_u = 2(\Sigma w)_{ij}\ \sigma^2_u$$

(15) $2\,E(v_{it+1}\, v_{jt}\, u_{t+1}\, u_t) = 2[E(v_{it+1} v_{jt}) \cdot \sigma_{u1}] = -2(\Sigma w)_{ij} \cdot \sigma_{u1}$

Hence

$$\mathrm{Var}\,(\theta) = \frac{1}{T}\, G^{*\prime} G^{*}\, \sigma_u^2 + \sigma_{u1}\, (\frac{1}{T}\, G^{*\prime}\, G^{*}_{-1} + \frac{1}{T}\, G^{*}_{-1}\, G^{*\prime}) + 2\, \Sigma_{ws}(\sigma^2{}_u - \sigma_{u1})$$

$$= [(\frac{1}{T}\, G'G) + \rho_{u1}(\frac{1}{T}\, G'G_{-1} + \frac{1}{T}\, GG'_{-1})]\sigma_u^2$$

and

$$T\,\mathrm{var}\,(\beta_1 - \beta) = (\Sigma_{\Delta x \Delta w} \Omega_1{}^{-1} \Sigma_{\Delta x \Delta w})^{-1} (M_{\Delta x \Delta w} (M_{\Delta w \Delta w} - 2\Sigma w)^{-1} \cdot \mathrm{var}(\theta)$$

$$\cdot\, (M_{\Delta w \Delta w} - 2\Sigma_w)^{-1}\, M^1_{\Delta x \Delta w}) (\Sigma_{\Delta x \Delta w}\, \Omega_1^{-1} \Sigma_{\Delta x \Delta w})^{-1}$$

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