THE ECONOMIC JOURNAL

JANUARY 1994

The Economic Journal, **104** (*January*), 1–19. Published by Blackwell Publishers, 108 Cowley Road, Oxford OX4 1JF, UK and 238 Main Street, Cambridge, MA 02142, USA.

CONSUMER CONFIDENCE AND RATIONAL EXPECTATIONS: ARE AGENTS' BELIEFS CONSISTENT WITH THE THEORY?*

Daron Acemoglu and Andrew Scott

Using UK data the REPIH is rejected due to the predictive content of consumer confidence, and not labour income or any other macroeconomic variable. We provide evidence suggesting that this cannot be explained by liquidity constraints and account for this finding in terms of precautionary saving. Extending the CCAPM model to allow a time varying conditional variance we find a high level of confidence is associated with greater optimism about the level of consumption but also a higher forecast variance. Allowing for time aggregation, the overidentifying restrictions implied by this model are accepted. We estimate a small but statistically significant intertemporal elasticity of substitution.

Mistakes in predicting UK consumption largely account for the failure of forecasters to predict the strength of economic expansion in the late 1980s and the sluggishness of recovery in the early 1990s. This failure came after a decade of relative success using equations based on the error correction approach, pioneered by Davidson *et al.* (1978). Whilst various respecifications to such equations have been suggested, commentators have placed increasing emphasis upon measures of consumer confidence as a means of monitoring consumption trends. Unlike the error correction approach, focusing on consumer confidence explicitly introduces a forward-looking element,[1] which appears to cohere well with the standard theoretical approach to consumption, the Rational Expectations Permanent Income Hypothesis (REPIH, see Hall, 1978).

This use of confidence indicators raises a number of empirical and conceptual issues which this paper seeks to address. The empirical issues are whether consumer confidence is informative of current or future consumption, and whether consumer confidence has any predictive power over and above standard macroeconomic variables. The conceptual issues relate to the fact that confidence indicators reflect consumers' beliefs. The REPIH implies strong restrictions on the stochastic behaviour of consumption, given agents' beliefs about the future, and we examine whether consumer confidence is consistent with these restrictions.

* The work on this paper was completed while the first author was at the London School of Economics. We wish to thank Charlie Bean, Angus Deaton, Tony Garrett, Stephen Hall, John Muellbauer, Steve Pischke, David Walton, Clifford Wymer, seminar participants at the Bank of England and the CEP and two anonymous referees for their help, advice and comments. The normal disclaimer applies.

[1] The ECM approach is not necessarily inconsistent with forward-looking theories of consumption (see Favero and Hendry, 1990), as they include current dated explanatory variables. Thus they can be interpreted as the reduced form of a REPIH approach to consumption.

Section I shows consumer confidence is a coincident indicator of consumption growth and that this finding is consistent with the REPIH, as confidence is useful in predicting income growth as well as changes in housing wealth, inflation, unemployment and real interest rates. This predictive ability is over and above the influence of other macroeconomic variables. However, Section II shows that consumer confidence also predicts consumption growth, which is inconsistent with the REPIH. Considerable evidence is displayed to show that the growth of labour income (or any other macroeconomic variable) does not predict future consumption growth, whilst confidence does. Thus on UK data the REPIH is rejected in a different direction from that of previous studies, which have been due to excess sensitivity with respect to income.

The finding that consumer confidence predicts consumption even when we condition on income suggests that the consumption function shifts over the cycle. This raises the issue of whether these shifts can be explained within a modified form of the REPIH or whether they are caused by animal spirits (as suggested in Blanchard (1993)). Section III examines dropping two different assumptions of the REPIH: perfect capital markets and certainty equivalence. In the absence of closed-form solutions we use *ad hoc* non-parametric and parametric formulations and find no evidence that the predictive role of confidence indicators is due to imperfect capital markets. In examining whether precautionary behaviour can explain this role, we extend the consumption CAPM model of Hansen and Singleton (1983) to allow for a time-varying conditional variance. We propose a general framework to test whether the predictive role of any variable for consumption growth is due to precautionary saving. Assuming a constant relative risk-aversion utility function, precautionary saving implies that consumption growth depends on the variance of consumption and interest rates. Our results suggest that high consumer confidence signals both higher consumption growth and a higher forecast variance. As a consequence of precautionary saving, consumption grows faster due to the higher forecast variance. Once allowance is made for time aggregation, we find no additional role for confidence above this precautionary effect. Our empirical work therefore leads us to believe that non-certainty equivalence is an important aspect of rejections of the REPIH, at least on UK data, and that the predictive ability of consumer confidence does not necessarily imply a role for animal spirits.

I. CONSUMER CONFIDENCE AS A COINCIDENT INDICATOR

The measure of consumer confidence that we use in this study is based on a survey, available from 1974, performed by Gallup and commissioned by the EEC. Every month Gallup surveys a nationally representative group of 2,000 adults over the age of 16. Consumers are asked twelve questions in all, although we focus on a subset of five questions most pertinent to consumption. The responses to questions are graded on a scale of -1 to $+1$, in steps of 0·5. The overall consumer confidence measure we use is a simple average of responses to five questions: 'How do you think general economic conditions changed over the past 12 months?', 'How do you think general economic conditions will

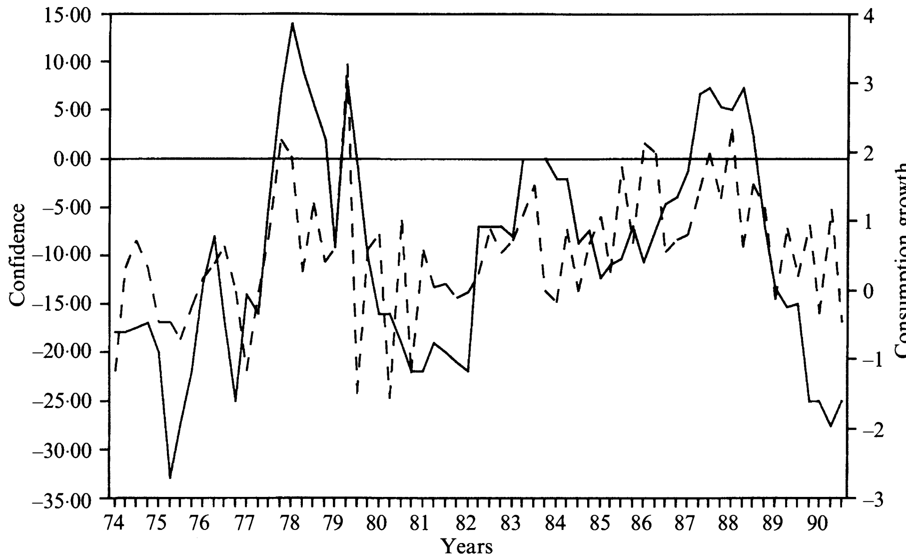

Fig. 1. Consumption growth (‐‐) and consumer confidence (——).

change over the next twelve months?', 'How do you think the state of your household finances has changed over the last 12 months?', 'How do you think the state of your household finances will change over the next 12 months?' and 'Is it a good time to make a major purchase?'.

Fig. 1 plots the quarterly average of the Gallup Consumer Confidence Indicator alongside the one-quarter change in the logarithm of non-durable consumption (a data appendix contains a description of all data used). The two series show a close correlation suggesting confidence may usefully play the role of a coincident indicator. Confidence indicates are available on a monthly basis and with a shorter publication lag than consumption data, and hence may serve as a good proxy for current unobserved consumption. Regressing non-durable consumption per capita on the current period confidence indicator (and a constant) yields a coefficient of 0·047 (with a t-statistic of 5·3) and an $\bar{R}^2$ of 0·29. There exist many potential explanations for this correlation. For instance, consumption may be driven by 'animal spirits', which might be directly reflected in the confidence indicator. Alternatively, forward-looking theories imply that confidence indicators may be a useful coincident indicator of consumption if they predict variables relevant to the consumers' planning problem. According to the REPIH (see Flavin, 1981) the change in consumption is proportional to the current innovation in future income expectations, i.e.

$$\Delta C_{t+1} = r \sum_{i=1}^{\infty} \left(\frac{1}{1+r}\right)^i (E_{t+1} - E_t)\, Y_{t+i}, \tag{1}$$

where r is a constant real interest rate, Y_{t+1} and C_{t+1} denote $t+1$ period labour income and consumption respectively. Therefore consumer confidence may be a coincident indicator if it summarises changes in agents' beliefs about future income. As two of the survey questions ask about agents' future expectations, the role of consumer confidence as a coincident indicator is potentially consistent with the REPIH.

To see whether the confidence indicator predicts income, we regressed the current change in the logarithm of labour income on lagged changes of itself, changes in the confidence indicator as well as the change in unemployment, inflation, real interest rates and changes in financial and housing wealth. All the explanatory variables were lagged one to four periods. The first column of Table 1 reports exclusion tests on each variable. Lagged income growth, the

Table 1

Granger Causality for Selected Macroeconomic Variables

	Dependent variable					
Deletion test on:	Income	Housing wealth	Financial wealth	Unemployment	Real interest rates	Inflation
Change in labour income	13·92	13·45	4·15	3·71	4·19	29·66
Change in housing wealth	10·96	17·68	3·79	16·25	1·26	8·25
Change in financial wealth	0·75	15·95	10·26	10·42	9·76	5·72
Change in unemployment	12·69	4·86	2·03	56·55	10·50	16·35
Real interest rate	5·36	4·46	2·37	30·34	39·55	5·34
Inflation	4·81	8·19	5·38	25·99	5·56	27·68
Confidence	12·22	14·42	0·75	17·85	9·36	12·61

Figure quoted is the LM statistic for the exclusion of lags one to four of the variable listed in the first column, the dependent variable is that listed in the column heading. Lags one to four of all variables listed in the first column were included in each regression. This test statistic is asymptotically distributed as χ^2_4, with 5% critical value 9·49 and 1% level 13·3. Sample period: 75q1–90q4. Estimation method: OLS.

confidence indicator and changes in unemployment and housing wealth are all found to predict labour income, other variables proving to be insignificant. Thus the REPIH can rationalise the role of consumer confidence as a coincident indicator for consumption as it predicts future income. The remaining columns of Table 1 repeat these causality tests for a wide range of macroeconomic variables. We find the confidence indicator useful in predicting innovations in all these variables, with the exception of changes in financial wealth (using an approximately exact sized test, as in Evans and Savin (1982), the test statistic for real interest rates is significant). This predictive role is present even when we condition upon a set of macroeconomic variables, suggesting that the confidence indicator reflects agents' private information.

However, whilst consumer confidence may reflect private information it is also heavily influenced by macroeconomic variables. Table 2 shows the results of regressing consumer confidence on the macroeconomic variables considered in Table 1. The final specification was arrived at using the 'general to specific' methodology, testing down from four lags and the contemporaneous value of each variable. This final specification has confidence depending positively upon its past value, negatively upon the current real interest rate and inflation

Table 2

Macroeconomic Determinants of Consumer Confidence

Variable	Coefficient	Robust t-statistic
Constant	6·063	2·280
Lagged confidence	0·664	7·869
Current real interest rate	−1·265	−5·566
Current inflation	−2·082	−3·068
Current change in housing wealth	0·831	1·921
$\bar{R}^2 = 0{\cdot}822$	S.E. 4·58	Serial correlation $\chi^2_4 = 0{\cdot}636$
Heteroscedasticity $\chi^2_1 = 0{\cdot}022$	Predictive failure test $\chi^2_{12} = 8{\cdot}043$	Parameter stability $\chi^2_6 = 7{\cdot}984$

Estimation method: OLS. Sample period: 1974q2 1990q4. Forecast and stability tests refer to 1988q1–1990q4. Dependent variable is confidence indicator.

Table 3

Confidence as a Leading Indicator for Consumption

Explanatory variable	Coefficient	R^2	S.E.
Confidence lagged, one period	0·042* (5·279)	0·440	0·686
Confidence lagged, two periods	0·028* (3·834)	0·292	0·777

Sample period: 1974q2–1990q3. t-statistics (in parentheses), calculated using White's standard errors. Constant and a dummy for the 1979 VAT change also included. Estimation method: OLS. * denotes rejection of REPIH at 5% significance level. Dependent variable is $\Delta \ln C_t$.

rate, and positively upon the current change in housing wealth. The confidence indicator is thus highly correlated with the current state of the economy and is a predictor of future economic strength, and can be thought of as a coincident indicator within the context of the REPIH.

II. CONSUMER CONFIDENCE AS A LEADING INDICATOR

Table 3 suggests confidence indicators are also a leading indicator for consumption. Whether we use confidence lagged one or two periods,[2] we find consumption growth to be predictable. Given our representative agent model, consumer confidence at time t is in the consumer's information set at time t. Hence the predictive content of consumer confidence violates the REPIH as outlined in equation (1), as neither the forward-looking nor backward-looking questions asked in the Gallup survey should contain any predictive ability for consumption growth.

Most previous studies reject the REPIH because consumption displays excess sensitivity with respect to lagged income (see, for UK data, Daly and Hadjimatheou, 1981, Muellbauer, 1983 and Campbell and Mankiw, 1989).

[2] As the confidence indicator is an average of a monthly series we need to make some allowance for time aggregation, which can lead to findings of predictability at a one-period lag, see Hall (1988).

Thus the results of Table 3 may reflect our finding that confidence indicators are a good proxy for anticipated income. Table 4 gives estimates of:

$$\Delta \ln C_t = \alpha + \beta d79_t + \lambda \Delta \ln Y_t + \epsilon_t, \quad (2)$$

where C_t denotes real *per capita* non-durable consumption[3] and Y_t denotes real *per capita* disposable labour income. The variable $d79_t$ denotes a dummy which takes on the value 1 in 1979q2, -1 in 1979q3 and zero elsewhere, and is included to account for pre-announced indirect tax changes in the 1979 budget. In testing for excess sensitivity we are interested in the anticipated component of income growth, and so we estimate the model using an instrumental variable-based estimator. To avoid the complications of time aggregation (Hall, 1988) we use instruments lagged two or more quarters. Time aggregation potentially introduces an MA(1) error into estimation, and if the REPIH is true we cannot arrive at correct inference using GLS (see Flood and Garber, 1981). To overcome this problem we use the alternative Aitken estimator suggested by Cumby *et al.* (1983); two-step two-stage least squares (2S2SLS). The fourth column of Table 4 quotes the coefficient and t-statistic on income and the final column the Sargan test for over-identifying restrictions on the instrument set. A significant coefficient on income suggests that consumption responds to anticipated income, whilst a significant value of the Sargan test shows that the instruments have a role in predicting consumption above that in predicting income growth. Either of these terms being significant is sufficient to reject the REPIH.

The first row of Table 4 shows that when only lagged income growth is used as instruments there is no evidence of excess sensitivity. When we add lags of confidence to the instruments, which significantly improves their quality, this finding is reversed but the Sargan statistic rejects the over-identifying restrictions. The third row of Table 4 reports estimates of

$$\Delta \ln C_t = \alpha + \beta d79_t + \lambda \Delta \ln Y_t + \gamma Conf_t + \epsilon_t, \quad (3)$$

where $Conf_t$ denotes the confidence indicator. We find no evidence for excess sensitivity with respect to income; instead it is the anticipated component of the confidence indicator which predicts future consumption growth and thus rejects the REPIH. Further, the inclusion of the confidence indicator leads to an insignificant Sargan statistic. Thus the results of Table 3 do not reflect the role of consumer confidence as a proxy for anticipated labour income. On the contrary, our results suggest that previous rejections of the REPIH on UK data were due to income proxying for consumer confidence.

Our results differ in two respects from those of Carroll *et al.* (1993), who perform a similar analysis using US data. First, the explanatory power of consumer confidence is considerably greater than for the United States, the largest $\bar{R}^2$ in their study being 0·17. Secondly, they find only weak evidence for

[3] Using the change in the logarithm as the dependent variable is not strictly speaking the appropriate way of testing the REPIH, unless we assume interest rates are deterministic as in Muellbauer (1983) and the utility function is CRRA. The advantage of our logarithmic specification is that it gives the best fit to the data, as well as providing a natural means of progressing to the variable-interest-rate model of Hansen and Singleton (1983).

Table 4
Testing for Excess Sensitivity

(1) Instrument set	(2) $\Delta \ln Y_t$	(3) r_t	(4) $\Delta \ln C_t$	(5) λ	(6) σ	(7) γ	(8) Sargan test
$\Delta \ln y(-2$ to $-4)$	0.001	—	0.019	−0.493 (−1.156)	—	—	0.080
$\Delta \ln y(-2$ to $-4)$, *Conf* $(-2$ to $-4)$	0.085	—	0.126	0.314* (3.666)	—	—	12.385*
$\Delta \ln y(-2$ to $-4)$, *Conf* $(-2$ to $-4)$	0.085	—	0.343	−0.175 (−1.012)	—	0.067* (4.551)	4.036
$\Delta \ln y(-2$ to $-4)$, $r(-2$ to $-4)$	0.023	0.563	0.034	−0.166 (−1.145)	0.082 (1.349)	—	4.691
$\Delta \ln y(-2$ to $-4)$, $r(-2$ to $-4)$, *Conf* $(-2$ to $-4)$	0.104	0.771	0.342	0.226* (2.238)	0.112 (3.241)	—	14.749*
$\Delta \ln y(-2$ to $-4)$, $r(-2$ to $-4)$, *Conf* $(-2$ to $-4)$	0.104	0.771	0.342	−0.137 (−1.121)	0.040 (1.032)	0.057* (3.577)	8.930
$r(-2$ to $-4)$, *Conf* $(-2$ to $-4)$	—	0.728	—	—	0.054 (1.974)	0.037* (4.429)	6.198

Column (1) lists the instrument set for each row. Column (2) shows the $\bar{R}^2$ from regressing the change in the logarithm of labour income on the instrument set, column (3) shows the $\bar{R}^2$ from regressing the interest rate on the instrument set. Column (4) shows the $\bar{R}^2$ from 2S2SLS estimates of consumption growth regressed on various combinations of confidence, interest rates, and income growth. Column (5) shows the coefficient on income growth, column (6) the coefficient on interest rates, and column (7) the coefficient on the confidence indicator. The final column is the value of the Sargan statistic for whether the instrument set satisfies the over-identifying restrictions. Sample period: 1975q1–1990q3. Estimation method 2S2SLS.

a predictive role of consumer confidence once they condition on income, and that lagged income is still significant even when conditioned on consumer confidence. The main difference between their study and ours is in the definitions of consumption used. However, we find exactly the same results as Tables 3 and 4 when we use their definitions of non-durable consumption excluding services, or consumption of services as the dependent variable.

Equation (1) is based on the assumption of a constant interest rate. In Hansen and Singleton's (1983) consumption CAPM model (CCAPM) with a stochastic rate of return, consumption growth may be predictable. As Tables 1 and 2 show the confidence indicator and interest rates to be correlated at various leads and lags, this might explain our excess sensitivity finding. The bottom half of Table 4 shows the results of estimating:

$$\Delta \ln C_t = \alpha + \beta \tilde{d}79_t + \lambda \Delta \ln Y_t + \sigma r_t + \epsilon_t, \tag{4}$$

where r_t is the real net UK Treasury Bill rate. According to Hansen and Singleton's model the coefficient on income growth should be zero, and σ should be the intertemporal elasticity of substitution. When only lagged interest rates and income growth are used as instruments (row 4), we find no evidence of excess sensitivity or of intertemporal substitution effects. The addition of consumer confidence to the instrument set leads to both terms becoming significant, but the Sargan test again suggests confidence indicators have additional predictive power. The penultimate row of Table 4 shows the result of adding confidence as an explanatory variable to (4). Again it is only the anticipated component of the confidence indicator which causes rejection. The last row of Table 4 shows the results of estimating (4), but instead of income growth we use the confidence indicator as an explanatory variable. In this case we find a small elasticity of intertemporal substitution, on the borderline of significance, and a strongly significant effect from the confidence indicator.[4]

Tables 1 and 2 reveal consumer confidence to be related to other macroeconomic variables aside from interest rates, and so the predictive content of the confidence indicator may reflect the influence of these variables. Table 5 shows 2S2SLS estimates of the relationship between the growth of

Table 5

General Excess Sensitivity Tests

Variable	Coefficient	t-statistic
Confidence indicator	0·039	2·727
Income growth	−0·113	−1·237
Real interest rate	−0·005	−0·124
Current inflation rate	−0·204	−1·672
Change in housing wealth	0·022	0·375
Sargan test = 9·174	$R^2 = 0·457$	

Estimation method: 2S2SLS. Dependent variable: *Per capita* consumption growth. Constant and 1979 VAT dummy also included. Sample period: 1975q1–1990q3. Instruments used were lags two to four of all explanatory variables.

consumption, the 1979 VAT dummy, income growth and all the explanatory variables in Table 2. Again we find only the anticipated component of the confidence indicator is useful in predicting consumption growth. A possible criticism of our results is that mentioned in Miron (1986): tests of the REPIH may be misleading on seasonally adjusted data due to the potential biases noted in Sims (1974) and Wallis (1974). We found that repeating the analysis of Tables 3–5 using unadjusted data and including seasonal dummies led to exactly the same findings; the REPIH is rejected due to consumer confidence and not disposable income or any other macroeconomic variable.

III. WHY DOES CONSUMER CONFIDENCE PREDICT CONSUMPTION GROWTH?

Whilst the previous section established a statistical connection between consumer confidence and future consumption growth, we have not examined in detail any conceptual explanations for this link. The fact that consumer confidence predicts consumption growth even when we condition upon other

[4] We performed similar regressions using all the variables from Table 1 which Granger caused our explanatory variables as instruments. We also used the explanatory variables of Davidson *et al.* (1978), e.g. ratio of non-durable consumption and labour income lagged two periods, and lags two to four of income growth, inflation and changes in income growth and inflation, as instruments. Once more the confidence term was significant and the over-identifying restrictions accepted.

macroeconomic variables implies that the consumption function is shifting over time. One possibility is that these shifts are the result of preference shocks, which are somehow reflected in the consumer confidence measures. However, while preference shocks would lead to consumption shifting even though income expectations remained constant, they could not explain the predictive role of consumer confidence. According to the REPIH only unanticipated preference shocks should affect consumption growth, hence taste shocks cannot rationalise the predictive role of the confidence indicator. Another possibility is that these shifts in the consumption function reflect 'animal spirits', and defy any rational economic explanation. Alternatively, it may be that we can explain this predictive role if we drop some of the assumptions underlying the standard REPIH. In this section we focus on two often suggested explanations of why the REPIH is rejected, and see if either can rationalise the predictive content of consumer confidence, and whether this explanation is consistent with the data.

(i) *Imperfect Capital Markets*

The assumption which has traditionally attracted the most attention as the cause of rejection of the REPIH is imperfect capital markets. Whilst the liquidity-constraints argument might appear most relevant to durables, a number of studies have found evidence that imperfect capital markets affect non-durable consumption (see, *inter alia*, Flavin, 1981, 1985; Hall and Mishkin, 1982; Zeldes, 1989*a*). The hypothesis of borrowing constraints is introduced in these studies to explain the fact that non-durable consumption growth is Granger caused by income growth. Assuming some persistence in income, a positive income shock today signals higher future income and an upwards revision to permanent income and consumption. However, under borrowing constraints the consumer is unable to borrow against this higher income, and so is unable to finance the increased consumption. As a consequence, when the higher income is realised next period consumption rises closer to its optimal level. The overall result is that on average consumption grows fastest when borrowing constraints operate (see Deaton, 1991). If, as Table 2 suggests, consumer confidence predicts income growth, a high value of confidence today signals higher future income. In other words, the confidence indicator is high because agents are responding positively to the two forward-looking questions regarding the economy and household finances. If consumers face borrowing constraints, they will be unable to respond to this revision in permanent income but will have to wait until the higher income materialises. As a result we might find that measures of consumer confidence have the ability to predict future consumption growth.

This particular interpretation of the confidence indicator raises a general difficulty for our attempt to rationalise the predictive role of consumer confidence. We cannot be sure exactly how consumers interpret the survey questions, nor what information their answers reveal. For instance, it may be that the consumer confidence indicator measures not the desire to consume but the agent's expectation of how feasible it is to consume. However, if this were the case we would find a negative coefficient on lagged consumer confidence. Consumer confidence would be low when agents were unable to borrow, but consumption growth next period would be high when the higher income was realised.

Explicit testing of this interpretation of the predictive ability of confidence indicators is complicated by the lack of closed-form solutions for consumption in the presence of liquidity constraints. We can however perform some *ad hoc* testing motivated by the theory. If the predictive role of confidence indicators is due to liquidity constraints, its coefficient should be lower when liquidity constraints are less binding. The late 1980s is widely seen as a period of financial innovation in the United Kingdom, with borrowing constraints declining in importance (i.e. Muellbauer and Murphy, 1989). Therefore if liquidity constraints explain the rejection of the REPIH, we should expect a lower coefficient on the confidence indicator in Table 2 over this sample period. The problem here is being able to model the severity of capital market imperfections. The simplest approach is to follow Muellbauer and Murphy (1989) and test the constancy of this parameter using various interactive dummy terms on the confidence indicator. The dummies we used were either 0/1 variables, taking the value 0 before a fixed date (either end 1986, 1987, 1988 or 1989) and 1 thereafter, or of the curvilinear form used in Muellbauer and Murphy (1989). For neither specification or any end date were we able to find significant changes in the coefficient.

Our failure to find significant effects from these interactive dummies may be due to the fact that our dummy variables are poor measures of financial innovation. An alternative way of identifying the severity of liquidity constraints is to use the difference between the personal sector borrowing and deposit rate. As noted by Hayashi (1987), the existence of such a mark-up is undeniable evidence in favour of imperfect capital markets. The higher this mark-up is, the smaller the consumer's budget set and the more limited are borrowing opportunities. If we use this mark-up as a proxy for the severity of liquidity constraint, the coefficient on the confidence indicator should vary positively with this mark-up, provided that the predictive power of consumer confidence is due to liquidity constraints. Fig. 2 plots this interest rate mark-up and Table 6 shows the results of estimating:

$$\Delta \ln C_t = \alpha + \beta r_t + \gamma(1 - \delta \exp^{-w_{t-1}})\, Conf_{t-1} + \epsilon_t \tag{5}$$

where w_t denotes the wedge between the loan rate and the real net Treasury Bill rate. For the confidence term to be explained solely by imperfect capital markets we require $\delta = 1$, as then confidence has no predictive content if capital markets are perfect ($w_t = 0$). If $\delta > 0$ but < 1 then liquidity constraints can explain some but not all of the predictive role of confidence indicators. The results provide no evidence in favour of imperfect capital markets; δ is not even significant at the 10% level. However, again it can be argued that focusing on the wedge does not successfully model the severity of liquidity constraints. In equilibrium this mark-up reflects both demand and supply effects, whereas we wish to measure only the latter.

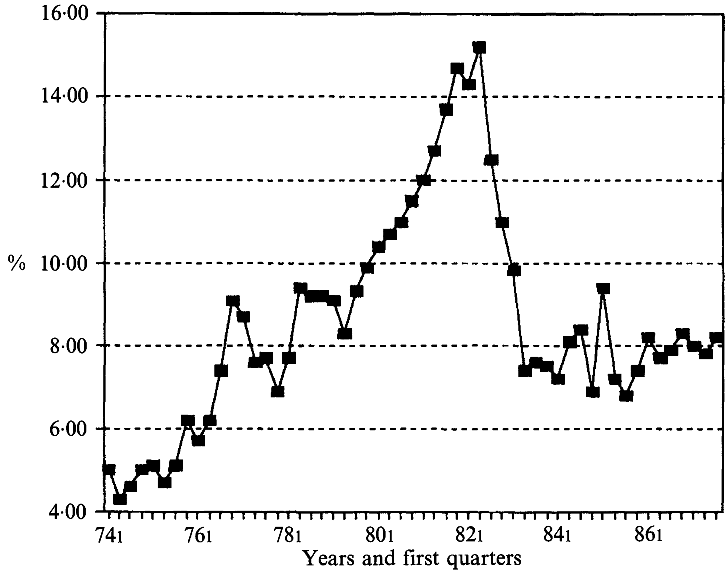

Fig. 2. Nominal borrowing and lending wedge.

Table 6

Testing for Imperfect Capital Markets

Parameter	Coefficient	t-statistic
α	0·855	3·612
β	0·023	0·523
γ	0·043	3·952
δ	0·113	0·244
$\bar{R}^2 = 0·430$	S.E. 0·704	

Estimation method: maximum likelihood. Algorithm: Davidson, Fletcher, Powell. t-statistics calculated using asymptotic standard errors. Equation estimated was:

$$\Delta \ln C_t = \alpha + \beta r_t + \gamma(1 - \delta e^{-w_{t-1}})\, Conf_{t-1} + \epsilon_t$$

A dummy was included for the 1979 VAT change as well. Sample period: 1975q1–1987q4.

Japelli and Pagano (1989) also cast doubt over whether for the United Kingdom a loan wedge is able to provide an adequate explanation for excess sensitivity with respect to income. They argue that in the United Kingdom liquidity constraints operate in the form of credit rationing, which is not reflected in the loan-deposit wedge. According to this argument, it is not surprising that our parametric specification (5) is unsuccessful. However, so long as we assume that the severity of credit rationing has not been constant over our sample we would expect some variability in the coefficient on consumer confidence, if the significance of confidence is due to imperfect capital markets. Therefore examining the recursive coefficient on consumer confidence is a robust and model-free way of examining whether borrowing constraints explain the predictive ability of the confidence indicators. Fig. 3

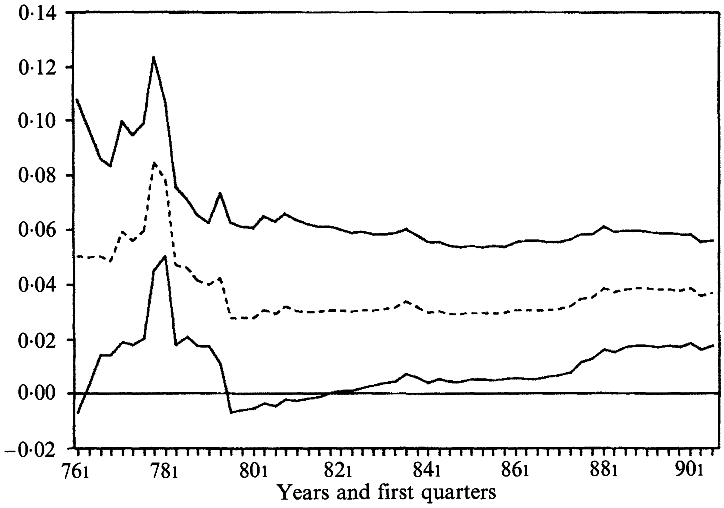

Fig. 3. Recursive coefficient on confidence. ---, Coefficient; ——, ±2 S.E.

plots the recursive coefficient on consumer confidence in Table 3. Except for some volatility in the early part of the sample, the coefficient is remarkably stable. This parameter constancy suggests that liquidity constraints are unlikely to explain the predictive content of consumer confidence.

(ii) *Precautionary Saving*

The REPIH is based on the assumption of certainty equivalence. The implications of income uncertainty for optimal consumption under non-certainty equivalence have been known for some time (e.g. Merton, 1969, Samuelson, 1969, Sandmo, 1974) but it is only recently that it has been proposed as a possible explanation for the rejection of the standard REPIH (e.g. Zeldes, 1989*a*). The basic idea underlying precautionary saving (abstracting from interest rate uncertainty) is that if agents are uncertain about their future income stream they will choose a lower level of current consumption, as they will wish to save in order to avoid low consumption as a result of low income states. However, their consumption profile must still satisfy the intertemporal budget constraint, and so as events unfold consumption will grow faster on average than under certainty equivalence. The slope of the consumption profile depends upon the variance of income, as this will determine how much saving the individual deems to be prudent. If this variance term is predictable, the REPIH would be rejected due to excess sensitivity.

As in the case of imperfect capital markets, the problem in testing precautionary saving lies in obtaining closed-form solutions. The work of Skinner (1988) and Caballero, 1990) has derived analogues to (1) (i.e. relating consumption growth to innovations in income) under various assumptions about preferences and the stochastic nature of interest rates and income. Pemberton (1993) outlines an alternative approach which essentially enables the consumer to approach the problem as if it were a two-period one. Rather than pursue these structural approaches we use a modified version of Hansen and Singleton's (1983) reduced-form approach. We assume a functional form for the variance of a linear combination of consumption growth and interest rates, which enables us to test whether the predictive content of confidence follows from precautionary behaviour.

Under the assumption that consumption growth and interest rates have a joint log-normal conditional distribution, and that agents' preferences can be represented by a CRRA utility function, we have

$$\begin{aligned} E_{t-1}\Delta \ln C_t &= (1/\rho)(E_{t-1}r_t-\delta)+\tfrac{1}{2}\rho\sigma^2_{t-1}, \\ \sigma^2_{t-1} &= \mathrm{var}_{t-1}[\Delta \ln C_t-(r_t/\rho)], \end{aligned} \tag{6}$$

where δ is the agents' rate of time preference and ρ is the coefficient of relative risk aversion. Hansen and Singleton (1983) examine the stochastic implications of (6) assuming a constant variance. Under these assumptions, the serial correlation properties of consumption growth correspond to those of $\{r_t\}$. More specifically, let the conditional expectation of interest rates be given by

$$E_{t-1}r_t = \mu_r+\mathbf{A}(\mathbf{L})\,\mathbf{Z}_{t-1}, \tag{7}$$

where $\mathbf{Z}_{t-1}$ is a $k\times 1$ column vector of predetermined variables and $\mathbf{A}(\mathbf{L})$ a $1\times k$ row vector of lag polynomials. According to (6)

$$\Delta \ln C_t = \frac{\mu_r}{\rho}-\frac{\delta}{\rho}+\tfrac{1}{2}\rho\sigma^2+\frac{1}{\rho}\mathbf{A}(\mathbf{L})\,\mathbf{Z}_{t-1}+\epsilon_t, \tag{8}$$

where ϵ_t is defined as $\Delta \ln C_t-E_{t-1}\Delta \ln C_t$.

However, if σ^2 is time-varying the over-identifying restrictions become more complex. Consider the system

$$\left.\begin{aligned} \Delta \ln C_t &= -\frac{\delta}{\rho}+\frac{1}{\rho}E_{t-1}r_t+\frac{\rho}{2}\sigma^2_{t-1}+\epsilon_t, \\ r_t &= \mu_r+\mathbf{A}(\mathbf{L})\,\mathbf{Z}_{t-1}+\mathbf{B}(\mathbf{L})\,\mathbf{X}_{t-1}+u_t, \\ \begin{pmatrix}\epsilon_t\\ u_t\end{pmatrix}|I_{t-1} &\sim N(\mathbf{o},\boldsymbol{\Sigma}+\mathbf{D}^*_{t-1}), \\ \mathbf{D}^*_{t-1} &= \begin{pmatrix}\mathbf{D}_{11}\mathbf{X}_{t-1} & \mathbf{D}_{12}\mathbf{X}_{t-1}\\ \mathbf{D}_{12}\mathbf{X}_{t-1} & \mathbf{D}_{22}\mathbf{X}_{t-1}\end{pmatrix}, \\ \boldsymbol{\Sigma} &= \begin{pmatrix}\sigma_{11} & \sigma_{12}\\ \sigma_{12} & \sigma_{22}\end{pmatrix}, \end{aligned}\right\} \tag{9}$$

where we have drawn a distinction between a vector of predetermined variables, $\mathbf{Z}_{t-1}$, which influence only the vector of means of the system, and an $m\times 1$ vector of variables, $\mathbf{X}_{t-1}$, which influence both the means and the covariance matrix. The vectors $\mathbf{D}_{ij}$, $i,j = 1,2$ are of order $1\times m$. We have assumed the variance–covariance term to depend linearly upon $\mathbf{X}_{t-1}$, although more general non-linear specifications are possible. Substituting the reduced form for interest rates into the consumption-growth equation and substituting for σ^2_{t-1} gives

$$\Delta \ln C_t = \mu^*_c + \frac{1}{\rho}\mathbf{A}(\mathbf{L})\,\mathbf{Z}_{t-1} + \left[\frac{1}{\rho}\mathbf{B}(\mathbf{L}) + \frac{\rho}{2}\left(\mathbf{D}_{11} + \frac{1}{\rho^2}\mathbf{D}_{22} - \frac{2}{\rho}\mathbf{D}_{12}\right)\right]\mathbf{X}_{t-1} + \epsilon_t. \tag{10}$$

Because our focus is on excess sensitivity, we relegate all the terms reflecting time preference into the constant. Testing for whether any variable predicts consumption growth due to precautionary behaviour requires estimating (9) as a system and testing the over-identifying restrictions implied by (10). Letting a_j denote the number of parameters in the jth column of $\mathbf{A}(\mathbf{L})$ and b_j the same concept for $\mathbf{B}(\mathbf{L})$, (10) imposes $\sum_1^k a_j + \sum_1^m b_j - 1$ over-identifying restrictions on (9). The results of Section II suggest that only consumer confidence has predictive power for consumption, and so we focus on the case where $m = 1$. In this case (10) implies very simple over-identifying restrictions, with each $\mathbf{D}_{ij}$ being a scalar. These restrictions involve both the serial correlation properties of r_t, through $\mathbf{A}(\mathbf{L})$ and $\mathbf{B}(\mathbf{L})$, as well as the covariance matrix of the system, through the $\mathbf{D}^*_{t-1}$ matrix. If the restrictions are rejected, not all the predictive content of the $\mathbf{X}_{t-1}$ variables can be explained by precautionary behaviour.

In choosing the appropriate variables for $\mathbf{X}_{t-1}$ and $\mathbf{Z}_{t-1}$ it is useful to relate the over-identifying restrictions implied by (10) to those of the instrumental variable approach of Section II. The restrictions of the $\mathbf{Z}_{t-1}$ terms are the same as a Sargan test for over-identifying restrictions when we instrument the interest rate term in (4) with $\mathbf{Z}_{t-1}$. Thus if the Sargan test is insignificant for a given instrument vector these are appropriate variables to use for the set $\mathbf{Z}_{t-1}$. However, if the addition of a variable to this instrument set leads to a significant value of the Sargan test, this variable should be included in $\mathbf{X}_{t-1}$. Obviously this way of allocating variables between $\mathbf{X}_{t-1}$ and $\mathbf{Z}_{t-1}$ means that the significance levels of the over-identifying restrictions should be interpreted with caution. If we had an actual series for the σ^2_{t-1} term, the tests of the over-identifying restrictions implied by (10) would be the Sargan test from regressing consumption on an interest rate and the variance term using $\mathbf{Z}_{t-1}$ and $\mathbf{X}_{t-1}$ as instruments. However, in the absence of a series for σ^2_{t-1} we cannot test the precautionary hypothesis using instrumental variables, but need to resort to a systems estimator which models σ^2_{t-1} directly.

The general approach to precautionary saving outlined above is similar in spirit to the work of Carroll (1992), who also suggests that rejections of the standard REPIH are due to variations in the conditional variance term in (6). However, our approach differs considerably from Carroll's in its implementation. Carroll selects a variable which he believes reflects agent's beliefs about income uncertainty (survey responses to questions regarding unemployment) and examines whether this predicts consumption conditional on income. In our approach we place a great deal of structure upon the unobserved variance term. An advantage of our approach is that we can test the over-identifying restrictions implied by precautionary behaviour, and not just examine statistical correlations. As a consequence, we do not have to rely as heavily on

Table 7
Testing for Precautionary Behaviour

(1) Explanatory variables	(2) $\bar{R}_r^2$	(3) $\bar{R}_c^2$	(4) σ	(5) Over-identifying restrictions	(6) Heteroscedasticity test
$\mathbf{Z}_{t-1}$ = {Real interest rate (−1 to −4), $\Delta \ln C$ (−1 to −4), confidence (−2 to −4)} $\mathbf{X}_{t-1}$ = {Confidence (−1)}	0·839	0·473	0·03 (1·75)	24·7** (11)	3·62
$\mathbf{Z}_{t-1}$ = {Real interest rates (−1 to −4), confidence (−2 to −4), change in financial wealth (−1 to −4), change in unemployment (−1 to −4)} $\mathbf{X}_{t-1}$ = {Confidence (−1)}	0·804	0·482	0·07 (2·21)	39·2** (15)	9·4**
$\mathbf{Z}_{t-1}$ = {Real interest rates (−2 to −4), confidence (−3 to −4), change in financial wealth (−2 to −4), change in unemployment (−2 to −4)} $\mathbf{X}_{t-1}$ = {Confidence (−2)}	0·723	0·376	0·05 (2·08)	16·1 (11)	8·6**

Column (2) quotes the R^2 for the restricted interest rate equation, column (3) for the restricted consumption growth equation. Column (4) is the estimate of the intertemporal elasticity of substitution. Column (5) quotes the LM test for the over-identifying restrictions implied by (10). This test statistic is distributed asymptotically as χ^2 with degrees of freedom given by the number in parentheses. The final column quotes the LM test for whether $\mathbf{D}_{t-1}^*$ is significant in the restricted version of (9). This test statistic is distributed asymptotically as χ^2 with 3 degrees of freedom and 5% critical value 7·81. Sample period: 1975q1–1990q3. ** denotes significant at the 5% level. The consumer confidence variable differs from that in previous tables as we normalise it to lie between 0 and 1 by adding 100 and dividing by 200.

the interpretation of the specific questions asked in the survey as Carroll (1992). As pointed out in Zeldes (1992), it is not clear to what extent the answers to the confidence survey reveal exactly the appropriate information for testing precautionary saving. However, in general any variable that predicts the variance of consumption and satisfies the restriction in (10) is informative about precautionary motives. In addition, in our case, there are plausible grounds to expect the confidence indicator to be informative about precautionary behaviour. On the other hand, a disadvantage of relying on the explicit testing of over-identifying restrictions is that, as it avoids interpreting the survey questions literally, it constitutes a black-box approach to precautionary behaviour. A further disadvantage is that any null hypothesis is a joint test of both our specification of the variance term and the precautionary version of the REPIH.

Table 7 shows the results of estimating (9), imposing the restrictions implied by (10), for various different choices of $\mathbf{Z}_{t-1}$ and $\mathbf{X}_{t-1}$. To ensure that our estimates of the variance of consumption and interest rates would be positive we decomposed $\mathbf{D}_{t-1}^*$ into $(\mathbf{D}'\mathbf{D})\, Conf_{t-1}$, where $\mathbf{D}$ is a symmetric 2×2 matrix. The first row shows the results when we use interest rates and consumption growth lagged one to four periods and consumer confidence lagged two to four periods as $\mathbf{Z}_{t-1}$, and the lagged confidence indicator as $\mathbf{X}_{t-1}$. The over-identifying restrictions implied by (10) are comprehensively rejected, and we can accept the restriction that the variance–covariance matrix is not time-varying. If we use the results of Table 1 to select our $\mathbf{Z}_{t-1}$ variables, the explanatory power improves but the over-identifying restrictions are still comfortably rejected, although the variance–covariance matrix now depends upon lagged confidence. All our estimates in Section II were based on instruments lagged two or more periods to avoid problems with time aggregation. The first two rows of Table 7 include confidence lagged one period, and this may explain the clear rejection of the over-identifying restrictions. To avoid this problem we include only variables lagged two or more periods in $\mathbf{Z}_{t-1}$ and $\mathbf{X}_{t-1}$. In this case we find that the over-identifying restrictions are accepted, whilst the dependence of the variance–covariance matrix upon consumer confidence remains. In other words, allowing for the time-aggregated nature of the confidence measure, we can account for all of the predictive ability of consumer confidence in terms of precautionary saving behaviour.

One limitation of our approach is that we have only considered the case where $m = 1$. Obviously, a proper specification search should be performed by adding additional variables to (9) in both $\mathbf{X}_{t-1}$ and $\mathbf{Z}_{t-1}$, although the dimension of the system soon becomes unwieldy. To ensure that we were not incorrectly excluding variables we performed LM tests for whether there was any remaining heteroscedasticity in (9) which could be modelled by any of the variables of Table 1, and in no case were these tests significant. In particular, Carroll (1992) suggests unemployment has a key role to play in determining precautionary behaviour. However, on UK data we could find no role for unemployment once we had conditioned upon consumer confidence. All three rows of Table 7 suggest that whilst agents do respond to changes in interest rates in setting their consumption profile, their intertemporal elasticity of substitution is numerically very small, similar to Hall's (1988) results on US data.

Examining the estimated $\mathbf{D}_{ij}$ in Table 7 we find the confidence indicator positively predicts the future variance of both consumption growth and interest rates and negatively predicts the covariance between them ($D_{11} = 0.107$, $D_{12} = -0.015$, $D_{22} = 0.32$). In other words, a high value of consumer confidence today suggests greater uncertainty about future values of interest rates and also about consumption growth. Agents respond to this uncertainty optimally by setting a lower level of consumption than they would under certainty equivalence, and thereafter consumption grows faster than it would under the standard REPIH. Further, this future growth is predictable in advance, as the conditional variance is correlated with lagged confidence. It may appear puzzling that a high value of consumer confidence predicts a high variance for consumption. The normal motivation of precautionary saving is that agents dislike uncertainty, in other words we would expect a low value of consumer confidence when the variance of consumption was high (e.g. as in Carroll (1992) where the variance term is linked to the unemployment rate). However, as high confidence predicts a high variance of consumption, it should also predict a high variance of labour income innovations as well as a high mean growth rate of income. Moreover, for precautionary behaviour to be reflected only in the confidence indicator, it must be only this variable which predicts the variability of the income stream. Regressing the squared residuals from the income process estimated in Table 1 on the squared fitted values gives a test statistic for heteroscedasticity of 5·568, which is distributed as χ_1^2. In other words, there is strong evidence that the variance of income is positively related to the magnitude of income growth. Regressing these squared residuals on a set of macroeconomic variables dated $t-2$ (to avoid time aggregation) the only variable which predicts the variance of income at the 5% level is consumer confidence, with a t-statistic of 3·81. Thus, when consumers become more optimistic about the future they also expect a larger variance for their forecast errors. Therefore, they only gradually revise upwards their consumption in response to a positive income shock in a manner consistent with precautionary saving.

The finding that a representative agent model with precautionary saving is accepted on UK data is in stark contrast to the results obtained for the United States. As explained by Deaton (1992), not only is the mean growth rate of US aggregate consumption positive and the mean real interest rate negative, but their relative variances are so small that (6) cannot possibly explain the data. This problem over the relative variances can be significantly lessened if we drop the assumption of a representative agent. In this case the variance term in (6) is no longer the variance of average consumption but the average variance of individual consumption, which is potentially much larger. However, on UK data the consumption growth and real interest rate series are not inconsistent with (6), even without introducing microeconomic heterogeneity. Over our sample period the mean growth rate of aggregate consumption is 0·54, whilst that of the quarterly real interest rate is 0·96%. Therefore, not only is there less work for the variance term in (6) to perform, but on UK data the variance terms are also larger than for the United States. The results in Table 7 also suggest that this variance term is procyclical, so that it is largest exactly when consumption is growing at its fastest, as our model requires.

IV. CONCLUSION

Confidence indicators are a useful coincident indicator for consumption over and above other macroeconomic variables, and this fact can be rationalised using the REPIH as they also predict future income growth. This ability to predict future income is additional to the predictive content of other macroeconomic variables, suggesting that confidence indicators reflect consumers' private information. However, the confidence indicator is also a leading indicator for consumption, contradicting the REPIH. Further, we find the REPIH is only rejected because of confidence indicators, and not because of excess sensitivity with respect to income or any other variable.

The final section of the paper sought to explain these findings, and in particular whether they shed any new light on why the REPIH is rejected. We find no evidence that the role of confidence indicators is caused by the existence of imperfect capital markets, due mainly to a stable coefficient on the confidence term. Allowing for time aggregation, and using a version of Hansen and Singleton's (1983) model in which the conditional variances of consumption growth and interest rates depend upon consumer confidence, we show that the predictive ability of confidence indicators is consistent with forward-looking behaviour and that we do not need to resort to animal spirits.

Massachusetts Institute of Technology

All Souls College, Oxford and Centre for Economic Performance, London School of Economics

Date of receipt of final typescript: July 1993

References

Blanchard, O. (1993). 'Consumption and the recession of 1990–91.' *American Economic Review Papers and Proceedings*, vol. 83, pp. 270–4.

Caballero, R. (1990). 'Consumption puzzles and precautionary savings.' *Journal of Monetary Economics*, vol. 25, pp. 1249–74.

Campbell, J. Y. and Mankiw, N. G. (1989). 'Consumption, income and interest rates: reinterpreting the time series evidence.' In *NBER Macroeconomics Annual 1989* (ed. O. Blanchard and S. Fischer), pp. 64–102. Massachusetts: MIT Press.

Carroll, C. D. (1992). 'The buffer-stock theory of saving: some macroeconomic evidence.' *Brookings Papers on Economic Activity*, vol. 2, pp. 61–135.

Carroll, C. D., Fuhrer, J. C. and Wilcox, D. W. (1993). 'Does consumer sentiment forecast household spending? If so, why?' *American Economic Review* (forthcoming).

Cumby, R. E., Huizinga, J. and Obstfeld, M. (1983). 'Two step two stage least squares estimation in models with rational expectations.' *Journal of Econometrics*, pp. 333–55.

Daly, V. and Hadjimatheou, G. (1981). 'Stochastic implications of the life cycle–permanent incomes hypothesis: evidence for the UK economy.' *Journal of Political Economy*, vol. 89, pp. 596–9.

Davidson, J., Hendry, D. F., Srba, F. and Yeo, S. (1978). 'Econometric modelling of the aggregate time-series relationship between consumers' expenditure and income in the United Kingdom.' Economic Journal, vol. 88, pp. 661–92.

Deaton, A. S. (1991). 'Savings and liquidity constraints.' *Econometrica*, pp. 1221–47.

Deaton, A. S. (1992). *Understanding Consumption.* Oxford: Oxford University Press.

Evans, G. B. A. and Savin, N. E. (1982). 'Conflict among the criteria revisited: the *W*, *LR* and *LM* tests.' *Econometrica*, vol. 50, pp. 737–48.

Favero, C. and Hendry, D. F. (1990). 'Testing the Lucas critique: a review.' Discussion Paper no. 220, Queen Mary and Westfield College, London.

Flavin, M. A. (1981). 'The adjustment of consumption to changing expectations about future income.' *Journal of Political Economy*, vol. 89, pp. 974–1009.

Flavin, M. A. (1985). 'Excess sensitivity of consumption to current income: liquidity constraints or myopia?' *Canadian Journal of Economics*, vol. 18, pp. 117–36.

Flood, R. and Garber, P. (1981). 'A pitfall in the estimation of models with rational expectations.' *Journal of Monetary Economics*, pp. 433–5.

Hall, R. E. (1978). 'Stochastic implications of the life cycle–permanent income hypothesis: theory and evidence.' *Journal of Political Economy*, vol. 86, pp. 971–87.

Hall, R. E. (1988). 'Intertemporal substitution in consumption.' *Journal of Political Economy*, pp. 339–57.

Hall, R. E. and Mishkin, F. S. (1982). 'The sensitivity of consumption to transitory income: estimates from panel data on households.' *Econometrica*, vol. 50, pp. 461–81.

Hansen, L. P. and Singleton, K. J. (1983). 'Stochastic consumption, risk aversion, and the temporal behaviour of asset returns.' *Journal of Political Economy*, vol. 91, pp. 249–65.

Hayashi, F. (1987). 'Tests for liquidity constraints: a critical survey and some new observations.' In *Advances in Econometrics, Fifth World Congress*, vol. II (ed. T. F. Bewley), pp. 91–120. Cambridge: Cambridge University Press.

Japelli, T. and Pagano, M. (1989). 'Consumption and capital market imperfections: an international comparison.' *American Economic Review*, vol. 79, pp. 1088–105.

Merton, R. C. (1969). 'Lifetime portfolio selection under uncertainty: the continuous time case.' *Review of Economics and Statistics*, vol. 57, pp. 247–57.

Miron, J. (1986). 'Seasonal fluctuations and the life cycle–permanent income model of consumption.' *Journal of Political Economy*, vol. 94, pp. 1258–79.

Muellbauer, J. N. (1983). 'Surprises in the consumption function.' ECONOMIC JOURNAL, vol. 93, pp. 34–50.

Muellbauer, J. N. and Murphy, A. (1989). 'Why has UK personal sector saving collapsed?' London: Credit Suisse First Boston.

Pemberton, J. (1993). 'Attainable non-optimality or unattainable optimality: a new approach to stochastic life cycle problems.' ECONOMIC JOURNAL, vol. 103, pp. 1–20.

Samuelson, P. A. (1969). 'Lifetime portfolio selection by dynamic stochastic programming.' *Review of Economics and Statistics*, vol. 51, pp. 239–46.

Sandmo, A. (1974). 'Two-period models of consumption decisions under uncertainty: a survey.' In *Allocation under Uncertainty: Equilibrium and Optimality* (ed. J. H. Dreze), pp. 24–35. New York: Macmillan.

Sims, C. A. (1974). 'Seasonality in regression.' *Journal of the American Statistical Association*, vol. 69, pp. 618–27.

Skinner, J. (1988). 'Risky income, life cycle consumption and precautionary savings.' *Journal of Monetary Economics*, vol. 22, pp. 237–55.

Wallis, K. F. (1974). 'Seasonal adjustment and relations between variables.' *Journal of the American Statistical Association*, vol. 69, pp. 18–31.

Zeldes, S. (1989*a*). 'Consumption and liquidity constraints: an empirical investigation.' *Journal of Political Economy*, vol. 97, pp. 305–46.

Zeldes, S. (1989*b*). 'Optimal consumption with stochastic income.' *Quarterly Journal of Economics*, vol. 104, pp. 275–98.

Zeldes, S. (1992). 'Comment on Carroll (1992).' *Brookings Papers on Economic Activity*, vol. 2, pp. 141–9.

DATA APPENDIX

All of the data are constructed from UK CSO National Accounts unless otherwise stated, and are available from *Economic Trends*. The majority of series are defined over the period 1955–91 and are quarterly and seasonally adjusted.

The consumption series is real non-durable consumption and is deflated by UK total population (*Annual Abstract*). A gross labour income series is constructed by adding income from employment, income from transfers and grants and self-employment income. Deducted from this are national insurance contributions and a proportion of total personal-sector tax payments. The proportion of labour tax paid is calculated as the share of gross labour income in total gross income (including income from dividends and rents). The resulting net series is then deflated by the overall consumer price index and the UK population series (*Annual Abstract*, linearly interpolated).

The financial wealth variable is obtained from CSO's *Financial Statistics* table S 14.5. The housing-wealth variable was calculated as the product of an average house price and the number of owner-occupied dwellings. The former was calculated using the Department of the Environment's index for house prices, scaled in terms of £ by using Table 10.9 of *Housing and Construction Statistics*, which gives the average value of properties mortgaged by building societies. The number of dwellings was interpolated from the annual series contained in *Housing and Construction Statistics*.

The interest rate series is the average discount rate for 91-day Treasury Bills published in CSO's *Financial Statistics*, table 13.8. The real interest rate was calculated by subtracting the annual inflation rate as measured by the consumer price deflator. This was converted into a net figure by multiplying by one minus the ratio of taxes on capital income paid to capital income received and then converted to a quarterly rate by dividing by four. The borrowing rate used in constructing a measure of the wedge was constructed as follows. Annual figures for total interest payments were drawn from the *Blue Book*, total mortgage payments (provided by the CSO) were deducted from this series to leave total payments on non-mortgage debt. This series was divided through by the stock of non-mortgage debt (*Financial Statistics*, table 14.4) to give an estimate of the interest rate paid on non-mortgage borrowing. The unemployment series used is that from table 21 in *Economic Trends Annual Supplement* (1992).