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The response of consumption to income
A cross-country investigation*

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In previous work we have argued that aggregate, post-war, United States data on consumption and income are well described by a model in which a fraction of income accrues to individuals who consume their current income rather than their permanent income. This fraction is estimated to be about 50%, indicating a substantial departure from the permanent income hypothesis. In this paper we ask whether the same model fits quarterly data from the United Kingdom over the period 1957–1988 and from Canada, France, Japan, and Sweden over the period 1972–1988. We also explore several generalizations of the basic model.

1. Introduction

During the last ten years, there has been a massive amount of research on aggregate consumption in the 'Euler-equation' tradition initiated by Robert Hall (1978). Recently, we have argued that many of the empirical findings in the literature can be explained by a simple model in which some fraction of income accrues to agents who follow the 'rule of thumb' of consuming their current income, and the remainder accrues to forward-looking optimizing agents. For aggregate U.S. data, the fraction of income going to rule-of-thumb consumers appears to be in the range 35% to 50% [Campbell and Mankiw (1989, 1990)].

The major goal of this paper is to see whether aggregate consumption in other countries can also be characterized in this way. Existing work has not answered this question definitively, for several reasons. Some studies use

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poor data: Campbell and Mankiw (1990), for example, lack international data on consumption of non-durables and services and use total consumption instead, while Jappelli and Pagano (1989) use annual data. Some studies can be criticized on econometric grounds: Jappelli and Pagano (1989) ignore various difficulties arising from unit roots and time aggregation of their data, while Bayoumi and Koujianou (1989) use a questionable seasonal adjustment procedure. Finally, the work of Davidson, Hendry, Srba and Yeo (1978), Davidson and Hendry (1981), and their successors uses a different theoretical framework and does not provide a direct answer to the question we ask here.

A second goal of this paper is to flesh out the interpretation of the rule-of-thumb model. It is tempting to argue that liquidity-constrained agents will consume their current income, so that our findings could result from liquidity constraints which bind only a fraction of the population.[1] But there is no reason to think that this fraction will be constant across countries. Liquidity constraints should be expected to affect more people in countries with poorly developed consumer credit markets, as Jappelli and Pagano (1989) point out. We ask whether the international pattern of our results is consistent with the liquidity constraints hypothesis.

It is often argued that liquidity constraints have declined in importance in many countries during the 1980's, as financial systems have been deregulated. We explore this idea by allowing the fraction of current-income consumers to vary over time in each country. Finally, we consider the claim that adjustment costs and 'near-rational' consumer behavior may have an important effect on aggregate consumption data [see Davidson, Hendry, Srba and Yeo (1978), Attfield, Demery and Duck (1989), Cochrane (1989) and Thaler (1990)].

The organization of the paper is as follows. In section 2 we review the modern literature on the permanent income hypothesis, emphasizing that the permanent income model fits the data well in certain dimensions and poorly in others. We then outline our alternative model and show how it can explain these disparate results. The section concludes with a discussion of estimation procedures for our model. Section 3 applies our methods to several international data sets. Section 4 interprets the results, and section 5 summarizes our findings.

2. The permanent income hypothesis and a simple alternative model

We begin this section by laying out the linear-quadratic intertemporal optimization problem used by Hall (1978) and Flavin (1981) to derive a permanent income consumption function. Although there are some well-

[1]Of course, a fully worked out model with liquidity constraints predicts more complicated behavior at both the individual and macro level [see for example Deaton (1991)]. The rule-of-thumb model can only be a first approximation to a model with liquidity constraints.

known difficulties with the quadratic formulation, it has the great advantage that it delivers a linear Euler equation. This can easily be combined with the linear budget constraint to derive a closed-form solution to the consumption problem: A 'consumption function'.

The consumer's problem is to choose consumption C_t to solve

$$\max U = \mathrm{E}_t \sum_{i=0}^{\infty} \beta^i u(C_{t+i}). \tag{1}$$

subject to

$$W_{t+1} = RW_t + YL_t - C_t, \tag{2}$$

and

$$\lim_{i\to\infty} \mathrm{E}_t[W_{t+i}/R^i] = 0. \tag{3}$$

Eq. (2) describes the evolution of wealth W_t over time; R is the constant gross interest rate, and YL_t is labor income. Eq. (3) is the constraint that prevents the consumer from running a Ponzi scheme, borrowing to finance an increase in consumption today and then borrowing forever to pay the interest on the debt. Eqs. (2) and (3) together imply that the budget constraint can be written as

$$W_t = (1/R) \sum_{i=0}^{\infty} (1/R)^i (\mathrm{E}_t C_{t+i} - \mathrm{E}_t YL_{t+i}). \tag{4}$$

The Euler equation for this intertemporal optimization problem is:

$$u'(C_t) = \beta R\, \mathrm{E}_t[u'(C_{t+1})], \tag{5}$$

and if $\beta R = 1$ and marginal utility is linear in consumption, then

$$C_t = \mathrm{E}_t C_{t+1}. \tag{6}$$

Substituting (6) into the budget constraint (4), one obtains the permanent income consumption function [Flavin (1981)],

$$C_t = (R-1)W_t + (R-1)/R \sum_{i=0}^{\infty} (1/R)^i \mathrm{E}_t YL_{t+i} \equiv YP_t. \tag{7}$$

Eq. (7) says that consumption equals the net interest rate times non-human wealth W_t, plus the net interest rate times human wealth where this is measured by the expected present value of future labor income. Permanent income YP_t is defined to equal the right-hand side of (7).

A useful alternative form of the consumption function, suggested by Campbell (1987), is obtained by defining total income $Y_t \equiv (R-1)W_t + YL_t$, and saving $S_t \equiv Y_t - C_t$. Eq. (7) implies that

$$S_t = -\sum_{i=0}^{\infty} (1/R)^i \mathrm{E}_t \Delta YL_{t+i}. \tag{8}$$

Saving takes place when current income is above permanent income and is expected to decline; in fact saving equals the expected discounted value of future declines in labor income. Eq. (8) also implies that if labor income is first-order integrated, saving is stationary and total income and consumption are cointegrated.

Finally, eqs. (7) and (2) imply that

$$\Delta C_t = (R-1)/R \sum_{i=0}^{\infty} (1/R)^i (\mathrm{E}_t - \mathrm{E}_{t-1}) \Delta YL_{t+i} = \Delta YP_t. \tag{9}$$

The change in consumption equals the change in permanent income, which equals the discounted value of revisions in expected future labor income.

This simple model can accommodate a number of changes to increase its realism. For example, we have written the budget constraint (2) without an error term for simplicity, but an error term can easily be added to represent 'unanticipated capital gains'.[2] The error has no effect on eqs. (5) through (8), and appears in eq. (9) with a coefficient $(R-1)$. Similarly, one can add a white noise error to the permanent income consumption function to represent 'transitory consumption'. This appears as a white noise error in eq. (8), and as a first-order moving average error in eq. (9).[3]

We can now summarize a large number of results on aggregate consumption behavior. Except where otherwise noted, these results apply to aggregate quarterly U.S. data in the post-war period.

In his seminal paper, Hall (1978) showed that other variables do not

[2]The error term must have an exogenous variance. One might wish to make the variance of the error term proportional to wealth, to capture the notion that ex post real interest rates are random. Unfortunately, once the variance of the error is related to a variable under the consumer's control, the certainty-equivalence property of the linear-quadratic framework is lost and the model becomes intractable. Below we present an alternative log-linear model which allows both ex post and ex ante real interest rates to vary through time.

[3]A number of other factors could cause a first-order moving average error to appear in (9); these include durability of goods, time aggregation, and white noise measurement error in the level of consumption. We discuss this further below. A serially correlated error could also be added to the consumption function, but there is a danger that the model becomes vacuous once a general error is allowed.

contribute greatly to consumption in forecasting future consumption. He concluded that eq. (6) is close to being satisfied, and in this sense the permanent income model is approximately true. Nelson (1987) redid Hall's work making some minor econometric improvements, and reached a similar conclusion.

Campbell (1987) studied savings behavior. He showed that saving Granger causes changes in labor income – as it must do if eq. (8) holds and if consumers have any information relevant for forecasting labor income beyond the history of labor income itself – and that the relation between saving and future labor income changes is negative. Saving seems to be quite highly correlated with an unrestricted forecast of the discounted value of changes in labor income, but is less variable than the unrestricted forecast.

Deaton (1987), West (1988), and Campbell and Deaton (1989) focused their attention on eq. (9). They argued that the time series process for labor income is highly persistent, so that the right-hand side of (9) is more variable than the change in labor income itself. Changes in consumption, on the other hand, are less variable than changes in labor income. These authors claimed that the smoothness of consumption is actually evidence against the permanent income hypothesis and not, as was traditionally supposed, evidence in favor of it.[4]

Finally, there is work in a rather different intellectual tradition by Davidson, Hendry, Srba and Yeo (1978) and Davidson and Hendry (1981). These authors estimated equations explaining U.K. consumption growth by contemporaneous and lagged income growth and one lagged level of consumption in relation to income. They interpreted the latter 'error-correction' variable as a measure of disequilibrium, which is slowly eliminated by partial adjustment of consumption towards its equilibrium value. Unfortunately, this work cannot be compared directly with the work on the permanent income hypothesis because the consumption equation includes contemporaneous income growth as well as lagged variables, so the significance of the lagged variables is not direct evidence for the forecastability of consumption growth.

2.1. *An alternative model*

In two earlier papers [Campbell and Mankiw (1989, 1990)], we have argued for a simple alternative model that can explain all these findings. The model sets aggregate consumption equal to a weighted average, with weights

[4]Allowing for unanticipated capital gains or transitory consumption does not help resolve the 'excess smoothness' puzzle. Unless these errors are strongly negatively correlated with the revision in permanent income, they will tend to make consumption noisy rather than smooth.

λ and $1-\lambda$, of current income and permanent income.[5] One can think of this average as resulting from a mixture of two types of agents in the population; current-income consumers receive a fraction λ of aggregate income, and permanent-income consumers receive a fraction $(1-\lambda)$. At this stage we will leave open the question of why any consumers should set their consumption equal to their current income, and will focus on the consequences of this assumption.

The 'λ model' implies that the consumption function is

$$C_t=\lambda Y_t+(1-\lambda)YP_t, \tag{10}$$

while saving is given by

$$S_t=-(1-\lambda)\sum_{i=0}^{\infty}(1/R)^i \mathrm{E}_t \Delta YL_{t+i} \tag{11}$$

and the change in consumption is

$$\Delta C_t=\lambda\Delta Y_t+(1-\lambda)\Delta YP_t=\lambda\Delta Y_t+(1-\lambda)\varepsilon_t, \tag{12}$$

where the notation ε_t is used to indicate that the change in permanent income is unforecastable.

These modifications to eqs. (7), (8) and (9) are simple, but they are sufficient to account for the various findings discussed above. First, the R^2 obtained when the change in consumption is regressed on all the relevant variables in consumers' information sets is

$$\begin{aligned} R_C^2 &= \mathrm{Var}(\mathrm{E}_{t-1}\Delta C_t)/\mathrm{Var}(\Delta C_t) \\ &= \frac{\lambda^2\,\mathrm{Var}(\mathrm{E}_{t-1}\Delta Y_t)}{\lambda^2\,\mathrm{Var}(\Delta Y_t)+(1-\lambda)^2\,\mathrm{Var}(\Delta YP_t)+2\lambda(1-\lambda)\mathrm{Cov}(\Delta Y_t,\Delta YP_t)}. \end{aligned} \tag{13}$$

If changes in permanent income and current income are not strongly negatively correlated, so that

$$(1-\lambda)^2\,\mathrm{Var}(\Delta YP_t)+2\lambda(1-\lambda)\,\mathrm{Cov}(\Delta Y_t,\Delta YP_t)\geqq 0,$$

then $R_C^2\leqq R_Y^2$, where R_Y^2 is the R^2 obtained when the change in disposable

[5]This model is a simpler version of Flavin's (1981) model. Flavin allowed consumption to depend on lagged as well as current income. (Below we consider allowing an extra lag of income growth to affect consumption growth.) See also Bayoumi and Koujianou (1989), Bean (1986), DeLong and Summers (1986), Hayashi (1982), Jappelli and Pagano (1989), and Summers (1982).

income is regressed on all relevant information. No matter what the value of λ, changes in consumption will not be highly forecastable if changes in disposable income are not highly forecastable. Hall's (1978) findings may be consistent with values of λ considerably different from zero.

Second, eq. (11) can account for Campbell's (1987) results. Under the λ model, saving is a multiple $(1-\lambda)$ of the value it would take under the permanent income hypothesis. It is therefore perfectly correlated with the optimal forecast declines in labor income, but is less variable.

Third, the λ model can explain the smoothness of changes in consumption. The variability of the change in consumption is

$$\lambda^2 \operatorname{Var}(\Delta Y_t)+(1-\lambda)^2 \operatorname{Var}(\Delta YP_t)+2\lambda(1-\lambda)\operatorname{Cov}(\Delta Y_t, \Delta YP_t),$$

and this can be less than either $\operatorname{Var}(\Delta Y_t)$ or $\operatorname{Var}(\Delta YP_t)$. Aggregate consumption is like a diversified portfolio, with reduced variance when $\operatorname{Cov}(\Delta Y_t, \Delta YP_t)$ is not too large.[6]

Finally, the λ model implies that disposable income and consumption obey an error-correction model. Both consumption growth and income growth are forecast by an error-correction term (lagged saving), but this term does not represent any kind of disequilibrium. It appears in the model because saving Granger causes income growth, and because when λ is nonzero any variable that forecasts income growth also forecasts consumption growth.

2.2. Logs and levels

Thus far our discussion has been couched in terms of levels of consumption. It is possible, however, to reformulate both the permanent income model and our alternative model in terms of log consumption and income. This has two advantages. First, the processes driving aggregate consumption and income seem to be log-linear rather than linear. Second, the log-linear model can accommodate time-varying ex ante real interest rates and random ex post real interest rates, whereas the discussion above assumed that the real interest rate is constant. The log-linear model has an offsetting disadvantage, however, which is that the interpretation of our coefficient λ in terms of the fraction of current income consumers can no longer be exact.

The log-linear model begins by assuming that the representative agent has power utility rather than quadratic utility:

$$u(C_t)=C_t^{1-\gamma}/(1-\gamma), \tag{14}$$

where γ is the coefficient of relative risk aversion, and $\sigma=1/\gamma$ is the

[6]Flavin (1988) makes this point in detail.

intertemporal elasticity of substitution. The first-order condition for optimal consumption choice is now

$$1=\mathrm{E}_t[\beta R_{t+1}(C_{t+1}/C_t)^{-\gamma}], \tag{15}$$

for any random asset return R_{t+1}. If we assume that asset returns and consumption are jointly conditionally lognormal and homoskedastic, then as Hansen and Singleton (1983) showed, the Euler equation simplifies to

$$\mathrm{E}_{t-1}\Delta c_t=\mu^*+\sigma\mathrm{E}_{t-1}r_t, \tag{16}$$

where μ^* is a constant and lower-case letters indicate the logs of variables.[7]

Our alternative model replaces this expression for expected consumption growth with

$$\mathrm{E}_{t-1}\Delta c_t=\lambda\mathrm{E}_{t-1}\Delta y_t+(1-\lambda)[\mu^*+\sigma\mathrm{E}_{t-1}r_t], \tag{17}$$

or equivalently

$$\Delta c_t=\mu+\lambda\Delta y_t+\theta r_t+\varepsilon_t, \tag{18}$$

where $\mu=(1-\lambda)\mu^*$, $\theta=(1-\lambda)\sigma$, and the error term ε_t is orthogonal to all variables known at time $t-1$ or earlier. If expected real interest rates are constant, (18) becomes

$$\Delta c_t=\mu+\lambda\Delta y_t+\varepsilon_t, \tag{19}$$

which is analogous to eq. (12) above.

It is also possible to derive a log-linear approximation to the consumer's budget constraint. Campbell and Mankiw (1989) combine this with the log-linear Euler equation to deliver an approximate log-linear consumption function. One can then restate all the results about the behavior of saving and the change in consumption in terms of the log consumption-income ratio and the growth rate of consumption. We do not repeat this analysis here, because the economic insights are much the same as for the linear version of the model.

2.3. Estimation

We now discuss how our model can be estimated and tested. We

[7]Eq. (16) can also be derived using a second-order Taylor expansion of the Euler equation (15). However for the intercept term to be constant, the joint distribution of asset returns and consumption must be homoskedastic.

concentrate on the log-linear versions of the model, eqs. (18) or (19), although the same methods can be used to estimate the linear version, eq. (12).[8]

Eqs. (18) and (19) cannot be estimated by Ordinary Least Squares (OLS), because the error term ε_t is orthogonal to lagged variables but not necessarily to Δy_t or r_t. The solution to this problem is to use Instrumental Variables (IV), where any lagged variables that help to forecast income and growth and real interest rates are appropriate instruments.

IV estimation of eq. (19) can be thought of as estimation of a restricted system of two equations. Consider regressing Δc_t and Δy_t onto a set of K instruments z_t. The regression system is

$$\Delta c_t = z_t\beta + \eta_{ct}, \qquad \Delta y_t = z_t\gamma + \eta_{yt}, \tag{20}$$

where β and γ are K-vectors of coefficients. The model (19) imposes that these two vectors are proportional to one another: $\beta = \lambda\gamma$. When lagged income and consumption growth rates are used as instruments, the system (20) is a VAR in differences. When lagged income and consumption growth and the consumption-income ratio are used as instruments, then (20) is an error-correction representation. In either case, (19) imposes a simple set of cross-equation restrictions.

These restrictions are of course testable. An LM test statistic can be formed by regressing the residual from the IV regression on the instruments. T times the R^2 from this regression, where T is the sample size, should have a χ^2 distribution with $(K-1)$ degrees of freedom if the λ model is well specified and the equation error is homoskedastic and serially uncorrelated. We use the R^2 from the residual regression as an informal diagnostic statistic, but we do not use the LM approach to test the model because it is hard to generalize this approach to handle conditional heteroskedasticity and autocorrelation in the equation error. Instead we use a Wald test, adding $K-1$ instruments as right-hand side variables in the IV regression, and testing the joint significance of the extra variables.[9]

2.4. Timing of the variables

When we apply the IV approach, we do not use instruments that are known only at time $t-1$. All the instruments we use are lagged at least two periods. There are several reasons for this conservative procedure.

[8]Campbell and Mankiw (1990) shows that for U.S. data, one obtains very similar results whether one works in levels or in logs.

[9]Engle and Kozicki (1990) give a general discussion of test procedures for restrictions of the form (19) on systems like (20).

First, the underlying model applies to consumption and income measured at points in time, whereas available data are time-averaged. It is well-known that time-averaging can produce spurious first-order auto- and cross-correlations, but higher-order correlations are not affected.[10] Second, it is possible that there is white noise measurement error in the levels of consumption and income, or that taste shocks create white noise 'transitory consumption'. This would lead to a first-order moving average [MA(1)] error in eq. (18), but the estimates of the parameters will be unaffected if instruments are lagged two or more periods.[11] Third, there may be some durability even in the goods labelled 'non-durables and services' in the National Income Accounts. For example, clothing and shoes are typically included in the category of non-durable consumption goods. With constant interest rates, durability leads to an MA(1) error structure in eq. (18) [Mankiw (1982)], but once again the parameter estimates are consistent if twice-lagged instruments are used. Finally, there may be delays in the publication of aggregate consumption and income statistics. If consumers know their own consumption and income, but do not know the aggregate numbers, then the aggregates are not legitimate instruments [Goodfriend (1986)]. We can mitigate this problem by using aggregate variables that are lagged one extra quarter.

In our empirical work, we also try varying the timing of the right-hand side variable in the IV regression. The basic model has Δy_t as the regressor, but it is possible that current-income consumers determine consumption by reference to y_{t-1} as well as y_t.[12] Therefore we try replacing Δy_t with a weighted average $\alpha\Delta y_t+(1-\alpha)\Delta y_{t-1}$. The parameter α measures the relative weight of current income growth as opposed to a one-period lag of income growth.

3. The λ model in international data

In Campbell and Mankiw (1989, 1990) we studied quarterly U.S. data on

[10]The original statement of this point is due to Working (1960). Christiano, Eichenbaum and Marshall (1991) argue that it is important for U.S. aggregate consumption data.

[11]Of course, serially correlated measurement error or transitory consumption would still lead to inconsistent estimates of the parameters. But once arbitrary errors of this sort are allowed, any behavior of the data can be explained.

[12]Carroll and Summers (1989) find that in the long run consumption moves closely with income. They argue that this can be reconciled with smaller high-frequency estimates of λ if myopic or current-income consumers respond to income with some delay. The model with lagged income is also in the spirit of Flavin (1981). The difference is that Flavin allowed enough lags for the model to be just-identified, whereas we allow a single lag and retain over-identifying restrictions. One other reason for including lagged income in the model is that there may be differences in the times at which income and consumption are measured within the quarter.

aggregate consumption of non-durables and services, in relation to personal disposable income, over the period 1953–1985. In this paper we compare the U.S. results with results from the U.K., where data are available over the period 1957–1988, and from Canada, France, Japan, and Sweden, where data are available over the period 1972–1988.[13] A detailed description of our data sources is given in the Appendix.

3.1. Seasonality

One issue that arises in studying international data is the treatment of seasonality. U.S. consumption and income data are only available in seasonally adjusted form, and this is true also for our Canadian and French data. But in the U.K. data are available in both seasonally adjusted and unadjusted form, while our Japanese data are unadjusted and our Swedish data are mixed (consumption is adjusted but disposable income is unadjusted).

The seasonally unadjusted data have a common seasonal in consumption and income, but this may reflect nothing more than consumers' preferences for market activity in one part of the year and non-market activity in another part. We certainly do not want to reject the permanent income hypothesis merely because there is a pre-Christmas boom in consumption and income. Accordingly it is essential to treat the unadjusted data in such a way that seasonal factors do not drive our results.

One possibility is to regress the data on seasonal dummies, but we found that this does not adequately remove seasonal patterns. In Japan, for example, the strength of seasonal variation has been growing strongly over time. For the unadjusted data we therefore work with annual growth rates, measured at a quarterly frequency.[14] This procedure is conservative in that it minimizes the chance that the permanent income hypothesis can be rejected by inappropriate handling of seasonality. (It also reduces the power of our test, because annual growth rates are harder to forecast than near-term quarterly growth rates. The results we report for Japan may reflect this loss of power.) For adjusted data, we work with quarterly growth rates. The availability of both adjusted and unadjusted data in the U.K. enables us to check the sensitivity of our results to the seasonal adjustment procedure.

[13] These sample periods are those available after reserving data for lags.

[14] This is also the procedure of Davidson, Hendry, Srba and Yeo (1978) and Davidson and Hendry (1981). Annual differencing introduces overlap into the error term of our equation, which we handle in the standard manner. Bayoumi and Koujianou (1989) deseasonalize their data using an exponential smoothing routine available in RATS, but we found that this routine introduces large spurious autocorrelations when it is applied to a random walk series.

Table 1

Forecasting disposable income growth using consumption and income.[a]

Country	Sample period	Seasonality	Lags 1–4 $\bar{R}_1^2$	Lags 1–4 $\bar{R}_2^2$	Lags 1–4 Granger causality	Lags 2–4 $\bar{R}_1^2$	Lags 2–4 $\bar{R}_2^2$	Lags 2–4 Granger causality
U.S.	53(1)–85(4)	SA	0.019 (0.001)	0.184 (0.000)	0.000	0.021 (0.006)	0.089 (0.000)	0.006
U.K.	57(2)–88(2)	SA	0.027 (0.171)	0.231 (0.000)	0.000	0.012 (0.181)	0.105 (0.007)	0.004
	57(2)–88(2)	NSA	−0.008 (0.982)	0.163 (0.000)	0.000	−0.007 (0.844)	0.124 (0.000)	0.000
Canada	72(1)–88(1)	SA	0.036 (0.103)	0.243 (0.000)	0.000	0.028 (0.287)	0.116 (0.000)	0.001
France	72(1)–88(1)	SA	0.035 (0.006)	0.104 (0.016)	0.048	0.047 (0.006)	0.128 (0.012)	0.023
Japan	72(2)–88(1)	NSA	0.014 (0.361)	0.064 (0.448)	0.283	0.030 (0.086)	0.046 (0.234)	0.419
Sweden	72(2)–88(1)	NSA	0.026 (0.085)	0.071 (0.081)	0.056	0.087 (0.004)	0.179 (0.020)	0.073

[a] $\bar{R}_1^2$ and $\bar{R}_2^2$ are the adjusted R^2 statistics from regressions of disposable income growth on its own lags, and on its own lags and lags of consumption growth and the log consumption–income ratio, respectively. The columns headed 'Lags 1–4' use lags 1–4 of income and consumption growth, and lag 1 of the log consumption–income ratio. The columns headed 'Lags 2–4' use lags 2–4 of income and consumption growth, and lag 2 of the log consumption–income ratio. The numbers in parentheses are the joint significance levels of the regressors. 'Granger causality' is the joint significance of the consumption variables where they are included.

Non-seasonally-adusted data are handled by using annual rather than quarterly growth rates. The annual growth rate of disposable income is regressed on its own lag, and on its own lag and one lag of annual consumption growth and the annual average log consumption–income ratio. In the columns headed 'Lags 2–4', the regressors are lagged by one extra quarter to remove overlap.

3.2. Forecasting income growth

One striking feature of the U.S. consumption and income data is that lagged income growth rates are poor forecasters of income growth. Much better income forecasts are obtainable when one adds lagged consumption growth rates and the lagged consumption-income ratio to the set of forecasting variables. This Granger causality from consumption to income should not be surprising, since it is an implication of the permanent income theory. But it does mean that it is important to use consumption variables as instruments when estimating the λ model on U.S. data.[15]

In table 1 we show that the data from other countries also have this feature. For each country the table shows the adjusted R^2 statistic from a regression of income growth on its own lags 1 through 4 or 2 through 4 (in

[15]Flavin (1981) used only income variables to forecast consumption.

the columns headed $\bar{R}_1^2$), together with the joint significance of these lags in parentheses.[16] Then the table shows the adjusted R^2 statistic from a regression of income growth on its own lags, the corresponding lags of consumption growth, and either lag 1 or 2 of the log consumption-income ratio.[17] This statistic appears in the columns headed $\bar{R}_2^2$, together with the joint significance of the forecasting variables in parentheses. Finally, the table reports significance levels for tests that the consumption variables Granger cause income growth.

The results in table 1 are striking. In the U.S., the U.K., and Canada, the adjusted R^2 statistic for forecasting income growth with lagged income growth is always below 5%, but this rises to anything from 9% to 24% when consumption variables are added to the regression. The consumption variables are jointly significant at the 1% level or better in every case. In France and Sweden, the results are similar but somewhat weaker statistically; the consumption variables Granger cause the income variables at only the 5% level (France) or the 10% level (Sweden). Japan is the exception to the general pattern; in Japan neither the income variables nor the consumption variables have any statistically significant forecasting power for income growth. This means that the coefficient λ is unidentified in Japanese data, since we cannot reject the hypothesis that current income and permanent income are equal.[18]

3.3. Forecasting consumption growth

In table 2 we estimate the λ model for our six countries. Given the results of table 1, we use as instruments the lags 2–4 of income and consumption growth, and lag 2 of the log consumption–income ratio. The first two columns of the table show the adjusted R^2 statistics for the regressions of consumption growth and disposable income growth, respectively, onto the instruments. The joint significance levels of the coefficients in these regressions are reported in parentheses; for the consumption growth regression, the joint significance of the coefficients is the test used by Hall (1978) to evaluate

[16]For seasonally unadjusted data, where annual differences are used, only one lag of income growth is included in the regression. This one annual lag covers the same span of calendar time as the four lags in the quarterly regression. In the columns headed 'Lags 1–4', the lagged annual difference runs from $t-4$ to $t-8$, whereas in the columns headed 'Lags 2–4' it runs from $t-5$ to $t-9$.

[17]For seasonally unadjusted data, the log consumption–income ratio is measured as an average over the period $t-4$ to $t-8$, or $t-5$ to $t-9$.

[18]Japan and Sweden, the two countries with the weakest Granger causality from consumption to income, also have the smallest effective data sets, since their data are seasonally unadjusted data over the period 1972–1988. This suggests that the Granger causality tests for these countries may have low power.

Table 2

Estimates of the λ model ($\Delta c_t = \mu + \lambda[\alpha \Delta y_t + (1-\alpha)\Delta y_{t-1}] + \varepsilon_t$).[a]

Country	Sample period	$\bar{R}_c^2$	$\bar{R}_y^2$	λ	α	Model test
U.S.	53(1)–85(4) (SA)	0.010 (0.230)	0.089 (0.000)	0.351 (0.117)	1.00 (NA)	−0.040 (0.859)
				0.363 (0.121)	0.929 (0.254)	−0.040 (0.772)
U.K.	57(2)–88(2) (SA)	0.083 (0.004)	0.105 (0.007)	0.203 (0.092)	1.00 (NA)	0.066 (0.017)
				0.372 (0.106)	0.306 (0.195)	−0.015 (0.179)
	57(2)–88(2) (NSA)	0.178 (0.000)	0.124 (0.000)	0.657 (0.138)	1.00 (NA)	0.073 (0.005)
				0.632 (0.144)	1.17 (0.135)	0.038 (0.332)
Canada	72(1)–88(1) (SA)	0.074 (0.067)	0.116 (0.000)	0.225 (0.107)	1.00 (NA)	0.009 (0.021)
				0.236 (0.111)	0.856 (0.360)	0.013 (0.996)
France	72(1)–88(1) (SA)	0.162 (0.000)	0.128 (0.012)	0.401 (0.208)	1.00 (NA)	0.111 (0.008)
				0.974 (0.346)	0.400 (0.189)	−0.022 (0.897)
Japan	72(2)–88(1) (NSA)	−0.045 (0.958)	0.046 (0.234)	0.035 (0.366)	1.00 (NA)	−0.044 (0.894)
				0.017 (0.439)	−2.01 (*****)	−0.046 (0.728)
Sweden	72(2)–88(1) (NSA)	0.074 (0.003)	0.179 (0.020)	0.357 (0.173)	1.00 (NA)	−0.038 (0.713)
				0.245 (0.428)	2.61 (6.08)	−0.048 (0.680)

[a]The instruments used in this table are Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, Δc_{t-2}, Δc_{t-3}, Δc_{t-4}, and $c_{t-2} - y_{t-2}$ for seasonally adjusted (SA) data, and $y_{t-5} - y_{t-9}$, $c_{t-5} - c_{t-9}$, and $\bar{c}_{t-5} - \bar{y}_{t-5}$ (where $\bar{x}$ denotes a four-quarter backwards moving average) for seasonally unadjusted (NSA) data. The first two columns show the adjusted R^2 statistics for regressions of consumption and income growth on the instruments, with significance levels in parentheses. The third and fourth columns give IV estimates of λ and α, with standard errors in parentheses. The last column gives the adjusted R^2 statistic for a regression of the IV residual on the instruments, with a p-value for a Wald test of the λ model in parentheses. All test statistics are heteroskedasticity- and autocorrelation-consistent. An entry of the form ***** indicates that a coefficient or standard error exceeded 10. An entry of the form (NA) indicates that the coefficient was restricted a priori rather than estimated.

the permanent income hypothesis.[19] All significance levels are adjusted for heteroskedasticity, so there is no deterministic relationship between the R^2 statistics and the significance levels, but as one would expect, higher R^2 statistics tend to be associated with stronger significance levels.

Lagged consumption and income variables have significant forecasting power for consumption growth in most of the countries in our sample. The greatest predictability occurs in France and in U.K. seasonally unadjusted data, where the adjusted R^2 statistics are over 15%; Canada, Sweden and U.K. seasonally adjusted data are somewhat less predictable, while the weakest results are for the U.S. and Japan.[20] The international pattern of predictability in consumption growth is quite close to the pattern of predictability in income growth, which is what one would expect if the λ model is true and the countries in our sample differ more in their income processes than in their values of λ. The extreme case is Japan, where disposable income growth is unforecastable, so if the λ model is right one should expect consumption growth also to be unforecastable.

3.4. The λ model

The right-hand side of table 2 reports IV estimates of the λ model. For each country we estimate both the basic model, in which $\lambda \Delta y_t$ appears on the right-hand side, and the augmented model, which has $\lambda[\alpha\Delta y_t + (1-\alpha)\Delta y_{t-1}]$ on the right-hand side. When $\alpha = 1$, the augmented model collapses to the basic model. The table gives coefficient estimates and asymptotic standard errors for the coefficients λ and α. The far right-hand column gives the adjusted R^2 statistic when the IV regression residual is regressed on the instruments, as well as the significance level from a Wald test of the hypothesis that the λ model describes the data.

The λ model is able to account for almost all the predictability of consumption growth in our sample of six countries. The over-identifying restrictions of the augmented λ model are never rejected, and in most countries the basic λ model does nearly as well. (In France and in seasonally adjusted U.K. data, however, the coefficient α is 0.3 or 0.4, and the augmented specification appears necessary if the λ model is to fit the data.)

The estimates of λ are both economically and statistically significant. Leaving aside Japan, where λ is unidentified, the estimates range from 0.2 in Canada, through 0.35 in Sweden and the U.S., to nearly 1.0 in France. The

[19] Hall worked in levels rather than logs, did not difference the data, and used once-lagged regressors. Nelson (1987) repeated his procedures using log differences and lagging the regressors two quarters as we do here.

[20] These results understate the overall predictability of consumption growth in the U.S. Table 3 below gives stronger evidence for forecastability when lagged real interest rates are used as instruments. Tables 4–6 give stronger evidence when lagged nominal interest rate variables are used. For a detailed discussion of the U.S. data, see Campbell and Mankiw (1989, 1990).

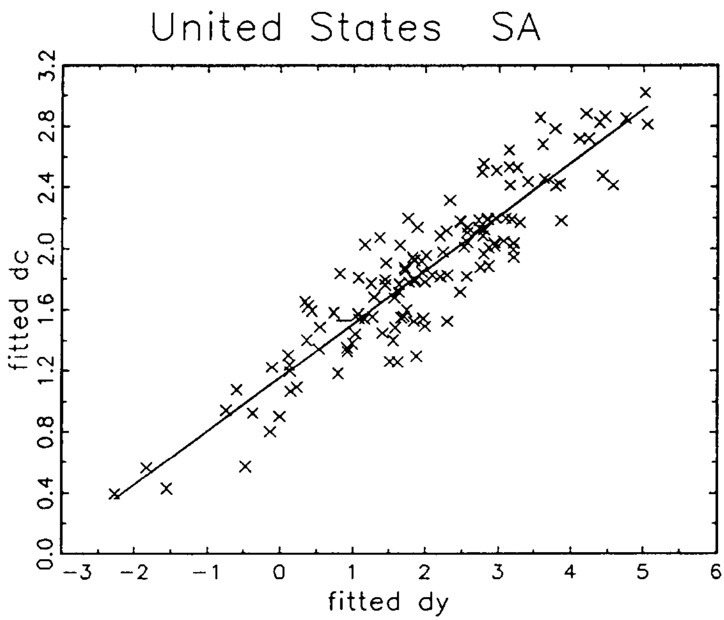

Fig. 1

results for the U.K. are somewhat sensitive to the seasonal adjustment procedure; λ is estimated at 0.35 in seasonally adjusted quarterly data, but is 0.65 when annual differences of seasonally unadjusted data are used.

In figs. 1 through 7, we give a visual representation of these results. The figures show the fitted values from regressions of consumption and income growth onto the table 2 instruments, along with the estimated IV regression line. Figs 2b and 5b, for U.K. seasonally adjusted and French data, replace the forecasts of income growth with forecasts of $\alpha\Delta y_t+(1-\alpha)\Delta y_{t-1}$, where α is set equal to the estimate in table 2. In every country except Japan, the forecasts of consumption growth and income growth have a well-defined linear relationship.[21]

Table 3 adds the short-term real interest rate to the specification of table 2, so that eq. (18) rather than eq. (19) is estimated. Lags 2 through 4 of ex post real interest rates are added to the set of instruments. This makes little difference to the results. The coefficient on the real interest rate is as often negative as positive, and it is never statistically significant when added to the

[21]This relationship is not simply due to the correlation between the ex post growth rates of consumption and income. Except in Japan, the ex ante correlation shown in the figures is stronger than the ex post correlation. See Campbell and Mankiw (1989) for a plot of ex post income and consumption growth rates in the U.S.

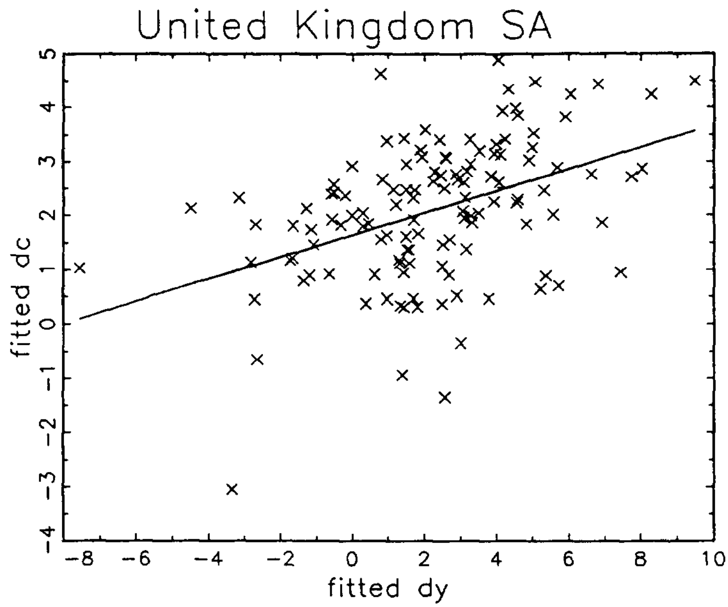
United Kingdom SA
fitted dc
fitted dy

Fig. 2a

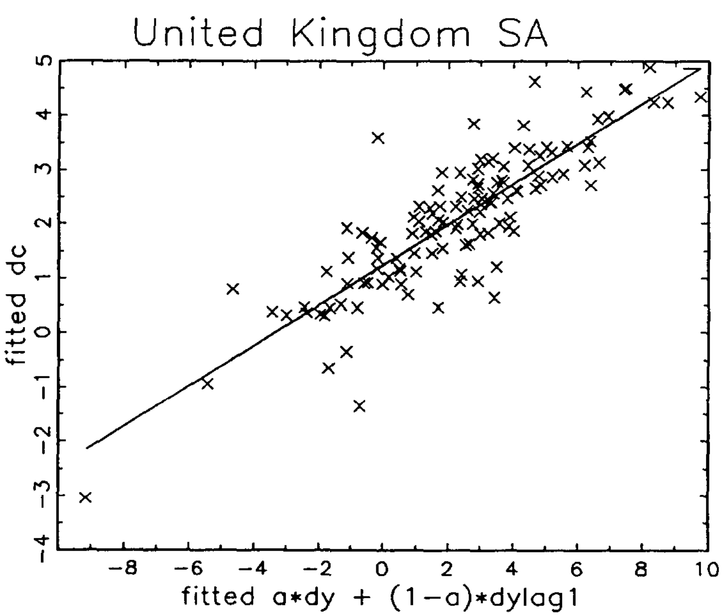
United Kingdom SA
fitted dc
fitted a*dy + (1−a)*dylag1

Fig. 2b

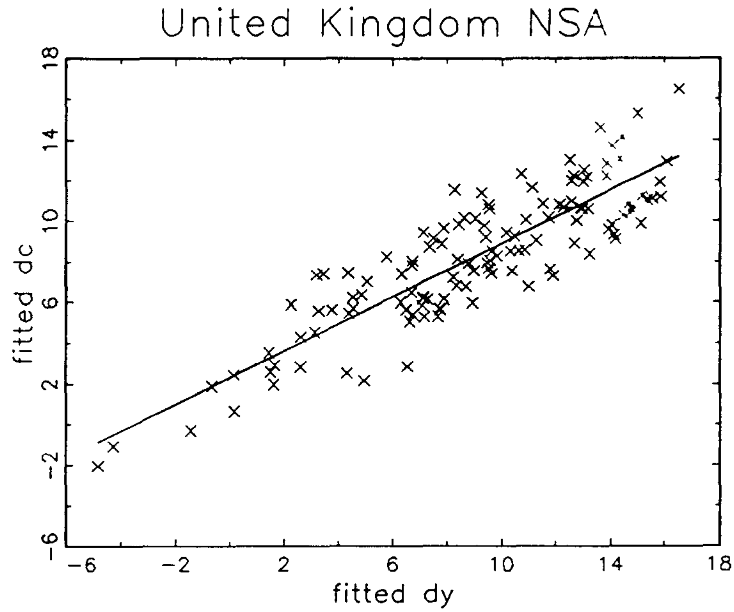

Fig. 3

preferred model from table 2. The estimates of λ tend to be somewhat lower in table 3 than in table 2, particularly in Canada and Sweden, but the overall pattern of the results is much the same.

3.5. Some specification tests

In table 4 we test the specification of the λ model by adding other variables to the right-hand side of the estimated equation.[22] We begin by adding the contemporaneous change in the nominal interest rate, Δi_t. Lags 2 through 4 of this variable are included in the instrument set in place of the lagged ex post real interest rates used in table 3. Wilcox (1989) has argued that changes in nominal interest rates may have a direct effect on consumption because indebted consumers, who face an upper limit on the ratio of their nominal debt service to their nominal income, must cut back consumption when nominal interest rates rise. However the fifth column of table 4

[22]See Campbell and Mankiw (1990) for other specification tests which add the stock of durables, labor supply, and government spending to the estimated equation for the U.S. These specification tests are motivated by a representative agent model in which the agent's utility is non-separable in consumption of non-durables and services and the level of these other variables.

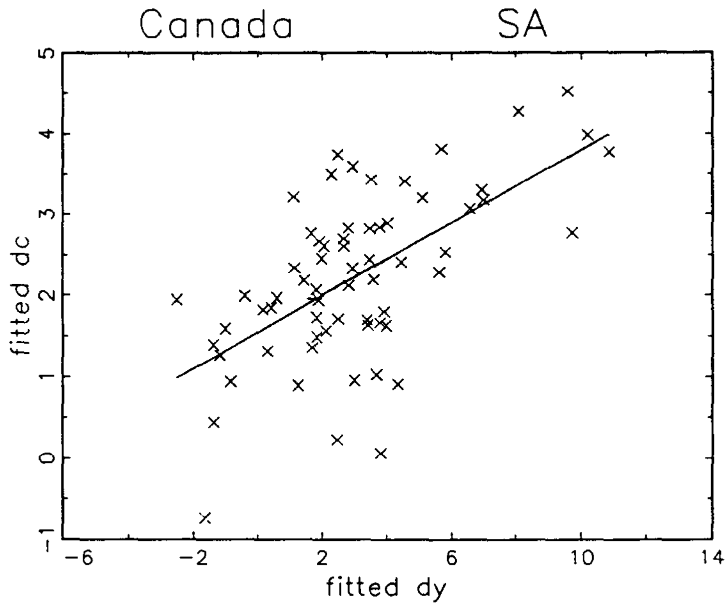

Fig. 4

shows that there is no statistically significant interest rate effect, once the effect of disposable income is allowed for.[23]

Table 4 also shows that in the U.S. and several other countries, the predictability of consumption and income growth is greater when lagged nominal interest rates are included in the instrument set. The instrument set makes little difference to the estimates of λ and α, which are generally quite similar to those reported in table 2; to save space, we do not report these estimates in the table.

The remaining columns of table 4 use the same instrument set, but add different variables to the basic specification. The sixth column adds the lagged change in consumption, Δc_{t-1}, to the equation. This variable would appear with a positive coefficient if there were important quadratic adjustment costs in non-durables and services consumption [Attfield, Demery and Duck (1989)]. In fact we find that the coefficient on Δc_{t-1} is often negative,

[23]Part of the reason for this finding may be that in several countries nominal interest rate changes are poorly forecast by our instruments. (This can be seen in the third column of table 4. It remains a problem even if we include a long-short yield spread in our instrument set; we do not report these results, as they are similar to those in table 4.) Also the interest rate used in this table is a 3-month Treasury bill rate, that is, a consumers' lending rate rather than a consumers' borrowing rate. It is possible that clearer evidence for an interest rate effect would be found if we used a borrowing rate.

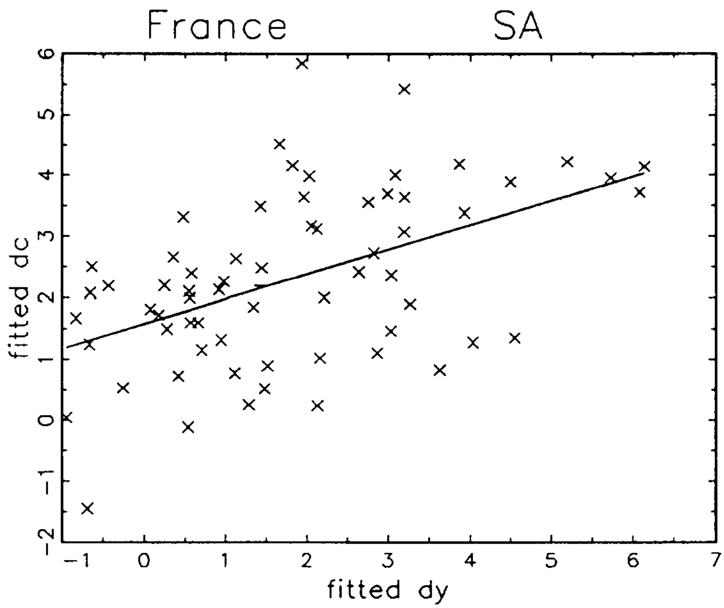
France
SA
fitted dc
fitted dy
6
5
4
3
2
1
0
−1
−2
−1
0
1
2
3
4
5
6
7

Fig. 5a

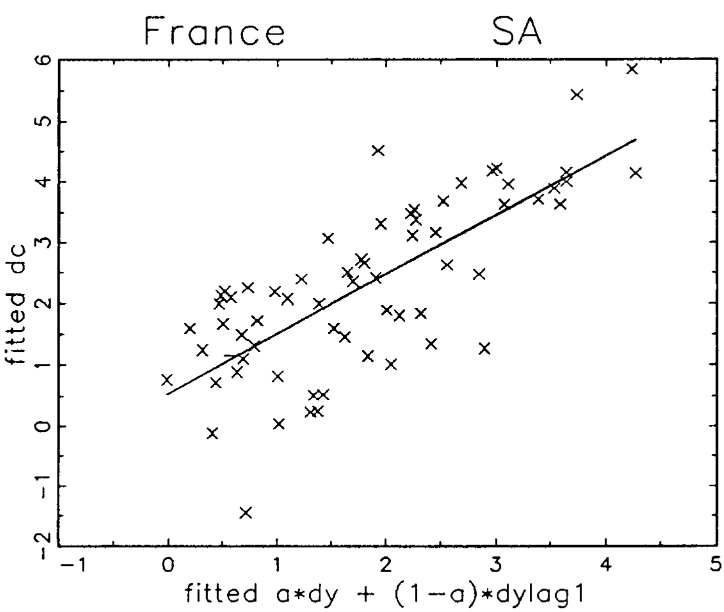
France
SA
fitted dc
fitted a*dy + (1−a)*dylag1
6
5
4
3
2
1
0
−1
−2
−1
0
1
2
3
4
5

Fig. 5b

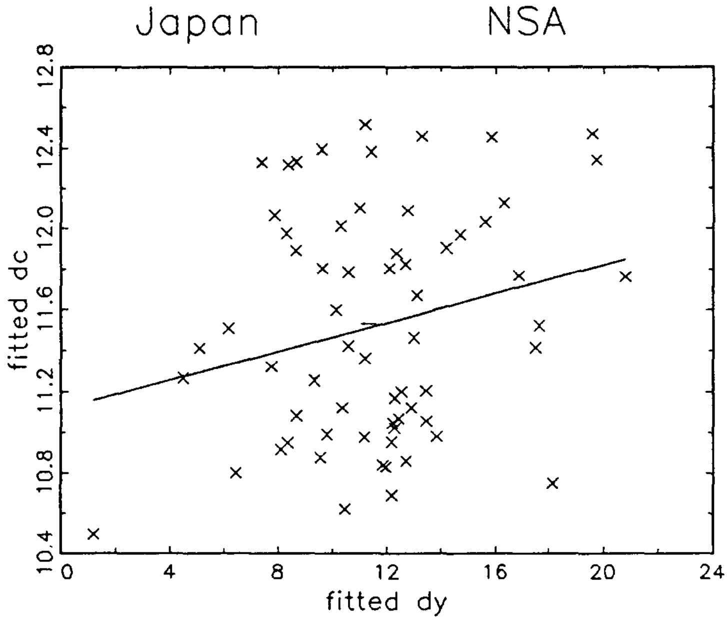

Fig. 6

and is never statistically significant when it is added to the preferred version of the λ model.[24]

The final column of table 4 adds a free error-correction term, $c_{t-1}-y_{t-1}$. This might appear with a negative sign in a disequilibrium model of consumption and income [see Davidson, Hendry, Srba and Yeo (1978), and Davidson and Hendry (1981)]. The error-correction term is never statistically significant, has no clear tendency to be positive or negative, and has little effect on the estimates of λ.[25]

4. Further interpretation of the λ model

In the previous section we showed that aggregate data on consumption and income in six countries are quite well described by a model in which a fraction λ of income accrues to current-income consumers, while a fraction

[24]The results are most favorable for partial adjustment in the U.S. data that were studied by Attfield, Demery and Duck (1989). Even here the partial adjustment term is insignificant.

[25]The error-correction term is largest (in absolute value) and most nearly significant in the seasonally unadjusted U.K. data that were studied by Davidson, Hendry, Srba and Yeo (1978), and Davidson and Hendry (1981). Even here it is not quite significant at the 5% level.

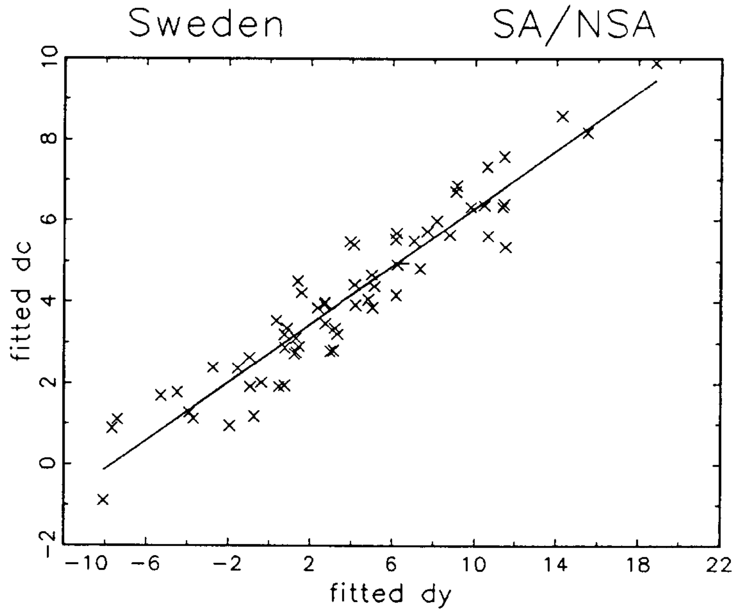

Fig. 7

$1-\lambda$ accrues to permanent-income consumers. In this section we try to interpret this result further. We first follow Cochrane (1989) in asking what welfare costs are implied by our model. Using a representative agent framework, Cochrane has showed that many alternatives to the permanent income hypothesis involve only trivial utility losses. We show that in the λ model, utility costs are indeed quite small if all agents share equally in aggregate consumption, but they are much larger if some agents are consuming current aggregate income.[26] This emphasizes the importance of research on the underlying structure that generates the λ model for aggregate consumption.

In the remainder of this section, we explore the idea that a positive estimate of λ reflects the importance of liquidity constraints. We compare the estimates of λ across countries, and ask whether countries with higher values of λ are also likely to have tighter liquidity constraints. Finally, we look to see whether λ has declined over our sample period; it is often argued that financial deregulation has relaxed liquidity constraints in the 1980's, and this might show up in a fall in the value of λ.

[26]Of course, utility costs would be much larger again if some agents were consuming their *own* current income.

Table 3
The λ model with real interest rates ($\Delta c_t = \mu + \lambda[\alpha \Delta y_t + (1-\alpha)\Delta y_{t-1}] + \theta r_t + \varepsilon_t$).[a]

Country	Sample period	$\bar{R}_c^2$	$\bar{R}_y^2$	$\bar{R}_r^2$	λ	α	θ	Model test
U.S.	53(1)–85(4)	0.031	0.090	0.489	0.385	1.00	0.004	−0.032
	(SA)	(0.045)	(0.000)	(0.000)	(0.104)	(NA)	(0.070)	(0.985)
					0.400	0.920	0.007	−0.032
					(0.114)	(0.232)	(0.070)	(0.830)
U.K.	57(2)–88(2)	0.069	0.121	0.321	0.017	1.00	−0.243	0.072
	(SA)	(0.007)	(0.004)	(0.000)	(0.062)	(NA)	(0.058)	(0.049)
					0.240	0.133	−0.121	−0.037
					(0.126)	(0.287)	(0.081)	(0.992)
	57(2)–88(2)	0.173	0.126	0.548	0.938	1.00	0.120	0.012
	(NSA)	(0.000)	(0.000)	(0.000)	(0.360)	(NA)	(0.139)	(0.045)
					1.17	1.17	0.235	−0.030
					(0.482)	(0.155)	(0.190)	(0.532)
Canada	72(1)–88(1)	0.067	0.083	0.552	0.085	1.00	−0.140	−0.012
	(SA)	(0.165)	(0.000)	(0.000)	(0.149)	(NA)	(0.158)	(1.000)
					0.056	1.39	−0.164	−0.017
					(0.246)	(3.76)	(0.221)	(0.985)
France	72(1)–88(1)	0.124	0.185	0.527	0.300	1.00	0.083	0.100
	(SA)	(0.000)	(0.000)	(0.000)	(0.164)	(NA)	(0.077)	(0.000)
					0.649	0.332	0.159	0.005
					(0.239)	(0.288)	(0.085)	(0.021)
Japan	72(2)–88(1)	−0.062	0.107	0.662	0.018	1.00	−0.022	−0.063
	(NSA)	(0.988)	(0.150)	(0.000)	(0.362)	(NA)	(0.113)	(0.923)
					−0.012	*****	−0.206	−0.068
					(0.340)	(*****)	(0.450)	(0.945)
Sweden	72(2)–88(1)	0.061	0.198	0.547	0.257	1.00	0.077	−0.050
	(NSA)	(0.006)	(0.000)	(0.000)	(0.123)	(NA)	(0.074)	(0.398)
					0.123	9.07	−0.062	−0.067
					(0.664)	(*****)	(0.527)	(0.741)

[a]The instruments used in this table are Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, Δc_{t-2}, Δc_{t-3}, Δc_{t-4}, $c_{t-2}-y_{t-2}$, r_{t-2}, r_{t-3}, and r_{t-4} for seasonally adjusted (SA) data, and $y_{t-5}-y_{t-9}$, $c_{t-5}-c_{t-9}$, $\bar{c}_{t-5}-\bar{y}_{t-5}$ and $\bar{r}_{t-5}$ (where $\bar{x}$ denotes a four-quarter backwards moving average) for seasonally unadjusted (NSA) data. The first three columns show the adjusted R^2 statistics for regressions of consumption growth, income growth, and the real interest rate on the instruments, with significance level in parentheses. The third, fourth, and fifth column give IV estimates of λ, α, and θ, with standard errors in parentheses. The last column gives the adjusted R^2 statistic for a regression of the IV residual on the instruments, with a p-value for a Wald test of the λ model in parentheses. All test statistics are heteroskedasticity- and autocorrelation-consistent. An entry of the form ***** indicates that a coefficient or standard error exceeded 10. An standard error entry (NA) indicates that the coefficient was restricted a priori rather than estimated.

Table 4
Some specification tests of the λ model ($\Delta c_t = \mu + \lambda[\alpha \Delta y_t + (1-\alpha)\Delta y_{t-1}] + \{\theta_1 \Delta i_t$ OR $\theta_2 \Delta c_{t-1}$ OR $\theta_3(c_{t-1} - y_{t-1})\} + \varepsilon_t$).[a]

Country	Sample period	$\bar{R}_c^2$	$\bar{R}_y^2$	$\bar{R}_i^2$	α	θ_1	θ_2	θ_3
U.S.	53(1)–85(4)	0.081	0.115	0.094	fixed	−0.134	0.222	−0.009
	(SA)	(0.001)	(0.000)	(0.412)		(0.556)	(0.236)	(0.018)
					free	−0.103	0.402	−0.010
						(0.551)	(0.345)	(0.019)
U.K.	57(2)–88(2)	0.074	0.149	0.081	fixed	0.942	−0.054	0.016
	(SA)	(0.005)	(0.001)	(0.014)		(0.700)	(0.173)	(0.018)
					free	0.804	−0.119	0.001
						(0.698)	(0.154)	(0.017)
	57(2)–88(2)	0.230	0.153	0.117	fixed	−0.645	−0.082	−0.115
	(NSA)	(0.000)	(0.000)	(0.009)		(0.626)	(0.129)	(0.066)
					free	0.516	1.12	−0.121
						(1.20)	(0.466)	(0.071)
Canada	72(1)–88(1)	0.079	0.129	−0.005	fixed	1.05	0.075	0.037
	(SA)	(0.047)	(0.000)	(0.231)		(0.724)	(0.285)	(0.033)
					free	1.08	0.026	0.051
						(0.710)	(0.326)	(0.046)
France	72(1)–88(1)	0.157	0.092	0.096	fixed	0.997	−0.034	0.003
	(SA)	(0.000)	(0.002)	(0.000)		(0.487)	(0.168)	(0.015)
					free	0.588	−0.179	0.018
						(0.550)	(0.212)	(0.019)
Japan	72(2)–88(1)	0.073	0.052	0.464	fixed	0.025	0.023	−0.082
	(NSA)	(0.138)	(0.332)	(0.000)		(0.838)	(0.306)	(0.096)
					free	0.874	0.468	−0.283
						(1.30)	(0.485)	(0.235)
Sweden	72(2)–88(1)	0.065	0.191	0.159	fixed	−0.256	−0.215	−0.000
	(NSA)	(0.006)	(0.004)	(0.002)		(0.442)	(0.291)	(0.082)
					free	−0.562	−0.213	−0.005
						(0.799)	(0.278)	(0.087)

[a]The instruments used in this table are Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, Δc_{t-2}, Δc_{t-3}, Δc_{t-4}, $c_{t-2} - y_{t-2}$, Δi_{t-2}, Δi_{t-3}, and Δi_{t-4} for seasonally adjusted (SA) data, and $y_{t-5} - y_{t-9}$, $c_{t-5} - c_{t-9}$, $\bar{c}_{t-5} - \bar{y}_{t-5}$ and $i_{t-5} - i_{t-9}$ (where $\bar{x}$ denotes a four-quarter backwards moving average) for seasonally unadjusted (NSA) data. The first three columns show the adjusted R^2 statistics for regressions of consumption growth, income growth, and the change in the nominal interest rate on the instruments, with significance levels in parentheses. The fourth, fifth, and sixth columns give IV estimates of θ_1, θ_2, and θ_3, with standard errors in parentheses.

Table 5

Utility losses implied by the λ model.[a]

Country	Sample period	λ	α	$\sigma(y_t - c_t)$	Loss measure L_1	L_2	L_3
U.S.	53(1)–85(4)	0.351	–	0.024	0.008	0.023	0.066
	(SA)	0.363	0.929	0.024	0.009	0.025	0.068
U.K.	57(2)–88(2)	0.203	–	0.038	0.005	0.023	0.114
	(SA)	0.372	0.306	0.039	0.026	0.070	0.188
	57(2)–88(2)	0.657	–	0.037	0.254	0.387	0.589
	(NSA)	0.632	1.17	0.037	0.200	0.317	0.502
Canada	72(1)–88(1)	0.225	–	0.037	0.006	0.026	0.114
	(SA)	0.236	0.856	0.037	0.007	0.028	0.120
France	72(1)–88(1)	0.401	–	0.042	0.039	0.097	0.241
	(SA)	0.974	0.400	0.041	116	119	122
Japan	72(2)–88(1)	0.035	–	0.030	0.000	0.002	0.048
	(NSA)	0.017	−2.01	0.094	0.000	0.008	0.461
Sweden	72(2)–88(1)	0.357	–	0.021	0.007	0.019	0.053
	(NSA)	0.245	2.61	0.056	0.016	0.067	0.274

[a]This table is based on the estimates of table 2. The first two columns repeat the coefficient estimates from that table. The third column gives the standard deviation of $\alpha y_t + (1-\alpha)y_{t-1} - c_t$. The final three columns give three measures of utility losses implied by these numbers. All losses are expressed in percentage points, as a fraction of optimal consumption. Loss measure L_1 is the right-hand side of eq. (24) in the text (the utility loss to a representative agent). Loss measure L_3 is the right-hand side of eq. (25) in the text (the utility loss to a current income consumer). Loss measure L_2 is λ times L_3 (the average utility loss to two groups of agents consuming current income and permanent income respectively). All these measures set γ, the coefficient of relative risk aversion, equal to 1. Utility losses for other values of γ are obtained by multiplying by γ.

4.1. The λ model and optimization failure

Cochrane (1989) has argued that recent empirical work on aggregate consumption has identified deviations from permanent income consumption that have only second-order effects on utility. He estimates that the utility costs of such deviations in U.S. non-durable consumption are only a few 1982 cents per quarter [see also Thaler (1990)].

Cochrane does not consider the λ model explicitly, but it is straightforward to calculate utility losses in this model. Consider the optimization problem (1), and write the actual consumption level as C_t, and the optimal level as C_t^*. The utility of the actual consumption path is U, and the utility of the optimal path is U^*. A second-order Taylor expansion gives

$$U - U^* \approx \mathrm{E}_t \sum_{i=0}^{\infty} \beta^i u'(C^*_{t+i})(C_{t+i} - C^*_{t+i})$$

$$+ (1/2)\,\mathrm{E}_t \sum_{i=0}^{\infty} \beta^i u''(C^*_{t+i})(C_{t+i} - C^*_{t+i})^2. \tag{21}$$

The first sum on the right-hand side of (21) is zero for any feasible perturbation from the optimal consumption plan. Utility costs can therefore be approximated by considering the second sum on the right-hand side of (21). The expectation of the first element of this sum, divided by u' to convert into dollars and then divided by C^*_t to express the per-period utility loss as a fraction of optimal consumption, is

$$(1/2)\,\mathrm{E}\left[\frac{u''(C^*_t)(C_t - C^*_t)^2}{u'(C^*_t)C^*_t}\right],$$

$$= (1/2)\,\mathrm{E}\left[\frac{C^*_t u''(C^*_t)}{u'(C^*_t)} \frac{(C_t - C^*_t)^2}{(C^*_t)^2}\right],$$

$$= (1/2)\gamma\,\mathrm{Var}(C_t - C^*_t)/C^*_t \approx (1/2)\gamma\,\mathrm{Var}(c_t - c^*_t), \tag{22}$$

where the equalities in the last line assume that the agent has power utility with relative risk aversion γ [eq. (14)], and that the proportional difference between actual and optimal consumption can be well approximated by the log difference.

The utility costs implied by the λ model depend on whether there is a single representative agent who consumes $c_t = \lambda y_t + (1-\lambda)c^*_t$, or two heterogenous groups consuming y_t and c^*_t respectively. In the former case, the representative agent's consumption is

$$c_t - c^*_t = (\lambda/(1-\lambda))(y_t - c_t), \tag{23}$$

and the expected per period utility loss to the representative agent, as a fraction of optimal consumption, is

$$L_1 = (\gamma\lambda^2/2(1-\lambda)^2)\ \mathrm{Var}(c_t - y_t). \tag{24}$$

If there are two groups of agents, the utility loss to the current income consumers is instead

$$L_3 = (\gamma/2(1-\lambda)^2)\ \mathrm{Var}(c_t - y_t), \tag{25}$$

and the average utility loss across both groups of consumers, L_2, is λ times this. Particularly for smaller values of λ, there can be large differences between the utility loss measures because $L_1 = \lambda L_2 = \lambda^2 L_3$.

Table 7 reports utility loss measures corresponding to the λ estimates of table 2. The table expresses utility losses in percentage points, as a fraction of optimal consumption, for estimated values of λ and for relative risk aversion γ equal to one. (Utility costs for other values of γ can be obtained by multiplying by γ.)

In most rows of table 7, the utility costs of the λ model to a representative agent are very small, less than 0.05% of optimal consumption and often less than 0.01%. The exceptions are the U.K. (seasonally unadjusted data) and France, where large values of λ lead to large utility costs: 0.25% of consumption in the U.K. and greater than 100% for France! (Obviously the second-order Taylor approximation breaks down for such large deviations from the optimum.)

Even where representative agent utility costs are small, they increase considerably when one takes account of possible heterogeneity across agents. The values of loss measure L_3 are usually 5 to 10 times larger than those of loss measure L_1. They vary considerably across countries, from about 0.05% in Sweden and the U.S., through 0.1% in Canada, to 0.2% in U.K. seasonally adjusted and 0.5% in U.K. unadjusted data. If γ were larger, say 10 rather than 1, these cost measures would be scaled up proportionally to 0.5% in Sweden and the U.S., 1% in Canada, and 2% or 5% in the U.K. These are quite serious welfare losses.

Of course, all these calculations assume that consumers face no idiosyncratic income risk. If current-income consumers face liquidity constraints that force them to consume their own current income rather than aggregate current income, the the utility costs would be much larger.[27]

4.2. The international pattern of results

The estimates we have presented indicate that there are differences across countries in the effect of current income on consumption. A rough summary of the results is that consumption is affected least by current income in Canada, somewhat more in Sweden and the U.S., more again in the U.K., and most of all in France.[28] In Japan we cannot reject the hypothesis that current income and permanent income are equal, so we cannot identify the direct effect of current income on consumption.

[27] Deaton (1989) presents an explicit model of liquidity constraints which can allow for idiosyncratic risk.

[28] The relative ranking of countries is much the same whether one looks at the predictability of consumption growth, the estimates of the coefficient λ, or the utility loss measures L_2 in table 7.

This pattern of results is consistent with previous findings in the literature, even though previous authors have used different data, different econometric methods, or both. Campbell and Clarida (1988) compared Canadian and British data, and found much stronger evidence against the permanent income hypothesis in Britain. Jappelli and Pagano (1989) ranked Sweden, the U.S., and the U.K. in the same order that we do, although they also estimated λ to be about the same in Japan as in the U.K.

These results also match fairly well with evidence about the development of consumer credit markets in the various countries. Jappelli and Pagano (1989), for example, report that among the countries they study, the ratio of consumer debt to total consumption is largest in Sweden, smaller in the U.S., and smaller again in the U.K. While we do not have data on consumer debt outstanding in Canada and France, it seems likely that consumer credit markets are well developed in Canada, but have been less so for most of the sample period in France. Thus the international pattern of results is consistent with the notion that the value of λ in aggregate data is determined by the prevalence of liquidity constraints.

4.3. Has λ changed over time?

Most of the countries in our sample have undergone substantial deregulation of their financial systems during the 1980's. It is sometimes argued that financial deregulation has relaxed whatever liquidity constraints may have affected aggregate consumption in the past, so that consumption is now free to move with permanent income and the value of λ is close to zero. In a similar spirit, the decline in U.K. personal saving in the second half of the 1980's is sometimes attributed to consumers' expectations of strong income growth, combined with their new ability to borrow.

Of course, there may well be constraints on consumer borrowing that result from equilibrium behavior of financial institutions in a world of asymmetric information. The large literature on credit rationing shows that such constraints can often be optimal. Even if financial deregulation removes constraints initially, the effect may last only until consumer debt builds up to the point where equilibrium liquidity constraints start to bind.

In tables 6 and 7 we look to see whether there is any evidence of a decline in the value of λ over time. Table 6 allows λ to be a linear function of a time trend, while table 7 allows for a one-time shift in λ at the end of 1979. The coefficient λ_0 in table 6 is the mean value of λ over the sample period, while the coefficient λ_1 is the change in λ per quarter. (The table reports $100\lambda_1$, which is the cumulative change in λ over 25 years.) In table 7, the coefficient λ_0 is the pre-1980 value of λ, and the coefficient λ_1 is the change in the 1980's. For simplicity, we do not estimate α simultaneously with λ_0 and λ_1,

Table 6

The λ model with a trend in λ ($\Delta c_t = \mu + (\lambda_0 + \lambda_1 t)[\alpha \Delta y_t + (1-\alpha)\Delta y_{t-1}] + \varepsilon_t$).[a]

Country	Sample period	$\bar{R}_c^2$	$\bar{R}_y^2$	λ_0	$100\lambda_1$	α	Model test
U.S.	53(1)–85(4)	0.027	0.115	0.354	0.152	1.00	−0.069
	(SA)	(0.012)	(0.000)	(0.090)	(0.185)	(NA)	(0.703)
				0.363	0.155	0.929	−0.067
				(0.091)	(0.188)	(NA)	(0.578)
U.K.	57(2)–88(2)	0.138	0.089	0.192	0.176	1.00	0.110
	(SA)	(0.000)	(0.007)	(0.088)	(0.203)	(NA)	(0.000)
				0.339	0.374	0.306	0.033
				(0.081)	(0.183)	(NA)	(0.250)
	57(2)–88(2)	0.380	0.228	0.709	0.470	1.00	0.031
	(NSA)	(0.000)	(0.000)	(0.096)	(0.175)	(NA)	(0.004)
				0.669	0.457	1.17	−0.014
				(0.097)	(0.164)	(NA)	(0.056)
Canada	72(1)–88(1)	0.077	0.075	0.200	−0.024	1.00	0.009
	(SA)	(0.003)	(0.000)	(0.121)	(0.341)	(NA)	(0.706)
				0.236	−0.008	0.856	−0.004
				(0.127)	(0.338)	(NA)	(0.858)
France	72(1)–88(1)	0.286	0.356	0.287	−0.958	1.00	0.073
	(SA)	(0.000)	(0.000)	(0.159)	(0.492)	(NA)	(0.008)
				0.295	−0.697	0.400	0.221
				(0.242)	(0.664)	(NA)	(0.596)
Japan	72(2)–88(1)	0.314	0.376	0.272	−0.538	1.00	0.135
	(NSA)	(0.000)	(0.000)	(0.275)	(0.520)	(NA)	(0.226)
				−0.359	−0.974	−2.01	0.097
				(0.179)	(0.519)	(NA)	(0.028)
Sweden	72(2)–88(1)	0.090	0.259	0.375	0.816	1.00	−0.093
	(NSA)	(0.000)	(0.000)	(0.157)	(0.649)	(NA)	(0.986)
				0.137	0.315	2.61	−0.034
				(0.087)	(0.337)	(NA)	(0.167)

[a]This table allows λ to be a linear function of a time trend. The instruments used are those from table 2, and the product of the table 2 instruments with the trend. The trend is demeaned so the coefficient λ_0 is the average value of λ in the sample, and the coefficient λ_1 is the change per quarter. The table reports $100\lambda_1$, the cumulative change in λ over 25 years. The coefficient α is fixed, rather than estimated, at 1.00 and the value estimated in table 2.

Table 7

The λ model with a 1980 shift in λ $(\Delta c_t = \mu + (\lambda_0 + \lambda_1 D_{80})[\alpha \Delta y_t + (1-\alpha)\Delta y_{t-1}] + \varepsilon_t)$.[a]

Country	Sample period	$\bar{R}_c^2$	$\bar{R}_y^2$	λ_0	λ_1	α	Model test
U.S.	53(1)–85(4)	−0.012	0.098	0.399	−0.123	1.00	−0.084
	(SA)	(0.000)	(0.000)	(0.118)	(0.136)	(NA)	(0.602)
				0.409	−0.122	0.929	−0.082
				(0.121)	(0.136)	(NA)	(0.510)
U.K.	57(2)–88(2)	0.059	0.064	0.091	0.194	1.00	0.047
	(SA)	(0.000)	(0.000)	(0.085)	(0.184)	(NA)	(1.000)
				0.203	0.282	0.306	−0.022
				(0.091)	(0.133)	(NA)	(1.000)
	57(2)–88(2)	0.250	0.111	0.589	0.574	1.00	0.010
	(NSA)	(0.000)	(0.000)	(0.139)	(0.147)	(NA)	(0.704)
				0.555	0.592	1.17	−0.042
				(0.140)	(0.122)	(NA)	(0.988)
Canada	72(1)–88(1)	0.170	0.113	0.240	−0.032	1.00	0.057
	(SA)	(0.000)	(0.000)	(0.082)	(0.140)	(NA)	(0.000)
				0.275	−0.039	0.856	0.052
				(0.093)	(0.137)	(NA)	(0.000)
France	72(1)–88(1)	0.223	0.299	0.444	−0.491	1.00	0.044
	(SA)	(0.000)	(0.000)	(0.169)	(0.226)	(NA)	(0.960)
				0.475	−0.222	0.400	0.149
				(0.198)	(0.280)	(NA)	(0.003)
Japan	72(2)–88(1)	0.415	0.448	0.518	−0.236	1.00	0.057
	(NSA)	(0.000)	(0.000)	(0.087)	(0.220)	(NA)	(1.000)
				0.032	−0.464	−2.01	0.162
				(0.160)	(0.187)	(NA)	(1.000)
Sweden	72(2)–88(1)	0.123	0.274	0.248	0.254	1.00	−0.079
	(NSA)	(0.000)	(0.000)	(0.118)	(0.253)	(NA)	(0.100)
				0.114	0.223	2.61	−0.090
				(0.090)	(0.354)	(NA)	(0.216)

[a]This table allows λ to change at the first quarter of 1980. The instruments used are those from table 2, and the product of the table 2 instruments with a time dummy equal to one from 1980(1) onwards. The coefficient λ_0 is the pre-1980 value of λ, and the coefficient λ_1 is the change in the 1980's. The coefficient α is fixed, rather than estimated, at 1.00 and the value estimated in table 2.

but fix it either at one (the basic λ model) or at the value estimated in table 2.

The results of tables 6 and 7 do not support the idea that liquidity constraints have declined in importance over time. There is no clear tendency for λ_1 to be positive or negative across countries, and in most countries it is statistically insignificant. The major exception is that in the U.K., λ_1 is significantly positive. This result does not come entirely from a trend or dummy shift in the mean growth rates of consumption and income, since λ_1 remains positive (although less significant statistically) when one allows for change over time in the intercept μ as well as the slope coefficient λ.

These time-series results contrast with the cross-sectional pattern of results. Cross-sectionally, countries with better developed credit markets seem to have lower values of λ; yet the evolution of credit markets over the post-war period does not seem to have caused a detectable decline in λ. This phenomenon deserves further investigation. It is possible that other factors, such as the increase in European unemployment in the late 1970s and 1980s, have worked to offset the effects of financial deregulation on λ. It is also possible that our methods, applied to aggregate data, are simply not powerful enough to detect movements in λ over time.

5. Conclusions

In this paper we have argued that aggregate consumption responds not just to changes in permanent income, but also to changes in current income. The direct effect of current income on consumption is measured by a coefficient that we call λ, which is scaled to equal zero if all agents consume their permanent income and one if all agents consume their current income. Using quarterly data from the U.S., the U.K., Canada, France, Japan, and Sweden, our main findings are as follows:

(1) Variables that predict income growth also predict consumption growth. In one country, Japan, neither income growth nor consumption growth are predictable, so the coefficient λ is unidentified. In the other countries predictable income growth and predictable consumption growth move in proportion with one another (with some adjustment in the timing of income growth for France and the U.K.). The factor of proportionality is the coefficient λ.

(2) Our estimates of λ range from 0.2 in Canada to nearly 1.0 in France, with Sweden, the U.S., and the U.K. falling in between. These values roughly match with previous estimates, and countries with higher values

[29] We also note that our time-series results contrast with those reported by Bayoumi and Koujianou (1989). It is not entirely clear to us why our results are different from theirs. Part of the explanation may be our different handling of seasonally unadjusted data.

of λ are those that are often thought to have less well developed consumer credit markets.

(3) We find no evidence that real interest rates or changes in nominal interest rates have direct effects on consumption growth once the effect of current income is taken into account. We also find no evidence that partial adjustment or disequilibrium error-correction terms help in modelling aggregate consumption once the effect of current income is taken into account.

(4) The utility losses associated with our estimates of λ are generally very small if the economy is made up of identical agents, each consuming a share of aggregate consumption. But they become much larger if the economy contains heterogeneous agents, some of whom consume aggregate current income, while others consume aggregate permanent income.

(5) It has been suggested that financial deregulation might have lowered the importance of current income for consumption in recent years. But we find no evidence that the coefficient λ has declined during our sample period.

Data Appendix

This paper uses data from several separate sources.

(1) For the United States, we use the data from Campbell and Mankiw (1990). These data were obtained from the Data Resources, Inc. data bank, and ultimately from the National Income and Product Accounts. All series are seasonally adjusted. The data cover the period 1947(1)–1985(4), but we start our sample period in 1953(1) for comparability with other work in the field, and because the data have an extreme outlier in 1950(1). See Campbell and Mankiw (1990) for further discussion; the data are reproduced in table 1 of that paper.

(2) For the United Kingdom, our data come from Central Statistical Office, *Economic Trends Annual Supplement 1989*.[30] Consumption and income data are available both seasonally adjusted (SA) and not seasonally adjusted (NSA). We use real personal disposable income (series codes CFAH and CFAG for SA and NSA data, respectively), total real consumption expenditure (CAAB and CCBH), and consumption expenditure on durables (CCBW and CCBI). The data cover the period 1955(1)–1988(2).

To convert the data to per capita form, we use IMF annual population data available from Datastream. We assume that the population numbers

[30]We used the printed source rather than series available in machine-readable form from Datastream and SWURCC. Preliminary checking found some serious problems with the machine-readable data; in particular, it seems that the 'seasonally adjusted' data are in fact not adjusted before 1970!

apply to the middle of each year, and create a quarterly population series by log-linear interpolation.

(3) Our third data set was kindly provided by Tamin Bayoumi and is described in Bayoumi and Koujianou (1989). The data cover six countries (Canada, France, Japan, Sweden, the United Kingdom, and the United States) over the period 1970(1)–1988(1). Consumption data are taken from the OECD *Quarterly National Accounts*; they differ from the longer consumption series we use for the U.S. and the U.K. in that 'semi-durable consumption expenditure' on items such as clothing and footwear has been removed.[31] Disposable income data come from various national sources. All series are seasonally adjusted except for Swedish income, and Japanese consumption and income. We convert the data to per capita form in the same way as for the U.K. data described above.

(4) Nominal interest rates on 3-month government debt are taken from the London Share Price Database.

[31]Blinder and Deaton (1985) make a similar correction to U.S. consumption data. In Campbell and Mankiw (1990) we found that the Blinder–Deaton data gave very similar results to the more standard National Income and Product Accounts data.

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